The Effects of Sectoral Decline on the Employment Relationship

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Abstract

This paper investigates tenure effects on employee retention under varying labor market conditions. We develop simple models of the likelihood of employer default (through early dismissals) on delayed payment and specific human capital contracts, and the predicted tenure pattern in these defaults under adverse conditions. The empirical implications of the models are investigated using competing risks analyses of tenure effects on recall and new job acceptance applied to layoff unemployment spell data from Waves XV and XVI (1982-83) of the Panel Study of Income Dynamics. The results indicate that adverse conditions (sectoral employment decline and unemployment) significantly reduce the positive tenure effect on recall probabilities in the 1983 data. This result is consistent with firm default on delayed payment contracts, but may also reflect the effect of technological changes that lower the value of firm-specific investments.
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I. Introduction

During the 1970s and 1980s, and continuing into the 1990s, U.S. industry has undergone substantial restructuring, largely due to factors such as changing patterns of foreign competition and technological change. The resulting shifts in labor demand across industries and occupational groups have produced significant worker dislocation. Although sectoral reallocation accelerates during recessions, long-run secular changes also occurred over the period (Loungani and Rogerson 1989). The impact of these factors is widespread. "Industrial downsizing" has now become a common term in the popular lexicon. Early retirement has increased since 1970 through at least 1986, with evidence for demand-side incentives as a primary motivating factor (Ippolito 1990). Filing of age discrimination suits also increased significantly during the 1980s (Wall Street Journal, 1987); although this may be primarily due to changes in the political environment surrounding age discrimination enforcement, it could also be linked to industrial downsizing and early retirement.

These phenomena suggest the potentially changing nature of employment security during periods of economic change. Employers largely adjust to economic fluctuations through the use of temporary layoffs and subsequent rehires. For example, Feldstein (1975) found that approximately 2/3 of all layoffs in manufacturing end in rehires (recall). However, the patterns in these rehires -- particularly for older, senior workers -- may change depending on the nature and severity of secular or short-run economic changes. Evaluating tenure effects on recall incidence is critical to the assessment of how binding are labor market arrangements that afford job security to senior workers.

Various theories, largely based on implicit contracts, have been invoked in the economic literature to explain worker/firm attachment. We focus here on two models that have the sharpest implications for
employers' incentive to retain workers at different tenure levels. Delayed payment contracts (Lazear 1979, 1981) are designed to solve the worker motivation problem under costly monitoring by paying wages below marginal product to relatively junior workers and wages above marginal product to relatively senior workers.\footnote{Throughout this paper, the terms "junior" and "senior" will be used synonymously with "low tenure" and "high tenure."} Although firms have an \textit{ex post} incentive to dismiss senior workers under these circumstances, reputational constraints minimize default incidence by increasing future labor costs for defaulting firms.\footnote{Abowd and Ashenfelter (1981) provide evidence for such reputational costs. They find substantial compensating wage premia in industries with high layoff histories.} In contrast, in the specific human capital model (Becker 1962; Hashimoto 1981; Mincer 1974; Mincer and Jovanovic 1981; Oi 1962; Parsons 1972) the rents resulting from shared training investments reduce firms' incentives to terminate senior employees.\footnote{Other models of worker/firm attachment are based on job matching (Jovanovic 1979), self-selection (Salop and Salop 1976), and insurance (Harris and Holmstrom 1982). We do not investigate the empirical implications of these models. See Garen (1988) for an evaluation of the evidence on job matching, Curme and Kahn (1990) for evidence comparing delayed payment and insurance models, Ryoo (1990) for an attempt to empirically distinguish between human capital and job matching models and Mortensen (1988) for a theoretical comparison. Also, Beaudry and DiNardo (1991, 1992) test general risk-sharing contract models using micro data.}

These implications, however, might require qualification during either a temporary or protracted economic downturn. For example, sectoral decline may increase the net benefits of default for delayed payment firms, thereby potentially increasing default incidence.\footnote{This idea was mentioned by Lazear (1979, p. 1271) in a footnote: "... a firm with an unanticipated decrease in its horizon will be more likely to cheat on the workers. Thus, one may expect to find that the incidence of pension default and early termination is higher in declining industries." Klein et al. (1978) discuss the relative costs of reneging on implicit contracts for firms with different growth rates.} In the specific capital case, economic dislocation or decline may reduce the present value of remaining training rents. Specifically, technological shocks may depreciate existing skills, thereby dissipating the rents from prior training investments and reducing incentives to retain senior workers. Similarly, if technological change is continuously occurring, firms may not find it profitable to continue to train older workers, particularly
if new required skills are more complementary to newer vintages of formal schooling.\textsuperscript{5}

This paper assesses the stability of the relationship between tenure and permanent dismissals -- as reflected in recall from temporary layoff -- during periods of sectoral decline or unemployment. Section II presents simple expositional models, based on the delayed payments and specific human capital frameworks, of the effects of adverse economic conditions on the relationship between tenure and firms' incentives for employee retention. Section III describes that data that we use to test the models' implications. These data include layoff unemployment spell data from Waves XV and XVI (1982-83) of the Panel Study of Income Dynamics (PSID), CPS sectoral employment tabulations, and data on industry technological change from Jorgenson et al. (1987). We analyze the spell data using competing risks models of recall and new job acceptance, as described in Section IV. Section V presents the empirical results and discusses their implications in light of our specific human capital and delayed payment models, focusing on the extent to which tenure patterns in recall probabilities are more consistent with the predictions of delayed payments or specific capital models.\textsuperscript{6} Our analysis uncovers significant mitigation of the tenure effect on recall in declining and high unemployment sectors, which we interpret as being largely consistent with employer default on delayed payment contracts. Section VI provides additional discussion of the results and suggestions for future work.

II. Rents and Employee Retention

The delayed payments and specific human capital models both possess features that minimize firm default in the form of premature dismissals. Under delayed payment contracts the mechanism is

\textsuperscript{5}Mincer and Higuchi (1988) make a similar argument in regard to Japanese labor markets, where the incidence (and average age) of mandatory retirement is greater (lower) in sectors with higher productivity growth; such sectors are presumably undergoing rapid technological change and face concomitant retraining requirements.

\textsuperscript{6}Studies that have attempted to empirically evaluate delayed payments models versus alternatives include Lazear (1979), Medoff and Abraham (1981), Lazear and Moore (1984), Hutchens (1986, 1987), Curme and Kahn (1990), Cornwell, Dorsey, and Mehrzad (1991), and Clark and Ogawa (1992).
reputation costs suffered by defaulting firms; in the specific human capital case the mechanism is shared rents following from shared investment. Each of these mechanisms, however, may be affected by adverse economic conditions. For example, the net costs of default may decline in response to economic changes that reduce the firm's horizon or increase its discount rate, and the net rents from training investments may decline in response to technological changes that depreciate existing firm-specific skills. In this section we provide simple models that illustrate these considerations and provide empirical implications.

(i) Delayed Payments Contracts

Assume that a firm offers a stylized version of the delayed payments contract described in Lazear (1979). Letting \( W(t) \) and \( V(t) \) denote the paths of wages and the value of the worker's output over time (tenure level) \( t \), a delayed payment contract sets:

\[
\begin{align*}
W(t) &< V(t) \quad \text{for } t < t' \\
W(t) &= V(t) \quad \text{for } t = t' \\
W(t) &> V(t) \quad \text{for } t > t'
\end{align*}
\]

subject to the constraint that the present values of \( W(t) \) and \( V(t) \) are equalized:

\[
\int_0^T [W(\tau) - V(\tau)] e^{-r \tau} \, d\tau = 0 \tag{1}
\]

where \( T \) specifies the completed length of the contract — i.e., the mandatory retirement date. The crossover point, \( t' \), is determined by conditions described in Lazear (1979, 1981); for our purposes, it is sufficient that \( 0 < t' < T \). Paying wages below marginal product to junior workers implies that workers are posting a bond; because repayment is contingent on performance, the contract increases worker motivation.

Under these conditions, employers have an incentive to default and fire workers when \( t > t' \). In a perfect information world with infinitely lived firms, however, this behavior will be recognized in the
labor market and will cause reputational effects that increase labor costs and prevent cheating firms from offering delayed payment contracts. With infinitely lived firms, these costs may completely eliminate firm default incentives. Yet if information is imperfect, or if firms experience an unanticipated increase in their failure probability (i.e. a shorter expected time horizon), some firm malfeasance may take place (Lazear 1981).

Specifically, firms will default through early dismissal when default profits are positive -- i.e., when the present value of expropriable worker rents exceeds the present value of default costs:

\[ \pi_d(t, r, T') > 0 \text{ if } \int_t^{T} [W(r) - V'(r)] e^{-r(t - r)} dr > \int_t^{T'} C(r) e^{-r(t - r)} dr \]  

In (2), \( \pi_d \) represents profits from default, \( r \) is the discount rate, \( T \) denotes the time of planned retirement, and \( T' \) denotes the firm's horizon -- the expected time until sale of the firm's assets, or the time when costs of firm default go to zero, whichever is less. \( C(r) \) denotes costs to the firm from early dismissal of a worker with \( r \) years of tenure; these are reputational costs that may include higher required wages, more worker shirking, or lower morale.

Equilibrium delayed payment contracts will minimize the incidence of firm default. Changing economic conditions such as sectoral decline, however, may alter firm default incentives. Assuming such change, the questions that we ask are: (i) what effect will this have on the probability of firm default through permanent dismissals?; and (ii) if firm default occurs, will junior or senior tenure workers be

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7Cheating may take other forms, such as altering promised wage profiles. For empirical reasons, we focus on early dismissals. Lazear (1979, p. 1267) also focused on this form of cheating: "Firm cheating takes the form of promising a worker a stream \( W(t) \) from 0 to \( T \), but dismissing the honest worker at some \( t < T \) and depriving him of the promised wage stream."

8Lazear (1981) explicitly incorporates the expected costs of worker shirking and worker-initiated separations into the expression for net default benefits. Exclusion of these expressions does not change our basic conclusions and considerably simplifies the exposition. We have also simplified the expression for worker rents by ignoring severance pay and end-payments, such as pensions. Ignoring pensions may be innocuous for fully vested workers covered by defined contribution plans. Dismissal of workers prior to vesting, however, might be one form of firm default. Unfortunately, our data does not contain information on pensions or severance pay arrangements.
more at risk?; i.e., conditional on default, what will the tenure pattern in default be? Concerning (i), from equation (2) it follows that if firms are initially engaging in optimal default, then an increase in the net returns from default will lead to an increase in its incidence. Alternatively, if firms are initially in a no-default equilibrium, an increase in the profits from default will lead to a positive level of default for firms that are marginal with respect to the profits from default. Concerning (ii), the optimal tenure level at which to dismiss workers is that which maximizes the profits from default. Because we are examining the effects of unanticipated sectoral decline, firms do not choose the optimal time to default on workers evaluated at time $t=0$. Instead, if the firm decides to default, default will occur for workers at or near the tenure level that maximizes net returns, where revenue is evaluated from the current period until time of planned retirement for each worker, and costs are evaluated over the firm’s operating horizon.

Adverse economic conditions may be viewed in two alternative ways, either as a reduction in the firm’s expected horizon ($T'$), or a rise in the firm’s discount rate ($r$). Consider the qualitative effects of these changes on profits from default ($\pi_d$), based on inspection of (2). First, assume that $T'$ is reduced. If $T' > T$, this will cause the present value of default costs to fall, increasing the likelihood that $\pi_d$ will be positive for a delayed payment firm, and hence increasing the probability of default. If $T' < T$, then the present value of default revenue also falls, and the net effect on default profitability is indeterminant. This latter case is unlikely, however, because $T' < T$ implies that under normal economic conditions the firm expects to fail prior to worker mandatory retirement, a condition that makes delayed payment contracts non-enforceable. We therefore interpret a declining firm horizon as generally implying an increase in net default benefits.

Second, assume that $r$ rises; in the limit, if failure is imminent, then $r \rightarrow \infty$, so that only the

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9 The optimal default tenure level is a poorly defined concept unless all workers begin employment with the firm at the same age. If not, a given tenure level will indicate different lengths of time until planned retirement. Hutchens (1986) finds, however, that employers who are likely to offer delayed payment contracts are unlikely to hire older workers, which suggests some degree of uniformity in age at hire under delayed payments. Our empirical results in Section V are invariant to substituting age-at-hire for age.
current period matters. This lowers the present value of the costs of dismissal ceteris paribus, thereby increasing default profits. The present value of revenue from early dismissals also falls, however, and the change in default profits is indeterminant; default increases only if costs decline more than revenue.

An increase in \( r \), however, increases the optimal default tenure level, because \((W-V)\) tends to increase with tenure under delayed payment contracts (Lazear, 1983). This is especially true given the pressure on managers to produce short-run profits in many U.S. companies, especially as economic conditions deteriorate (see Medoff and Abraham, 1981).

A simple linear specification illustrates these points concerning optimal firm defaults. Revenue from default may be modeled by assuming linear wage and productivity schedules:

\[
W(t) = \alpha_0 + \alpha_1 t \quad \text{and} \quad V(t) = \beta_0 + \beta_1 t, \quad \text{with} \quad \alpha_0 < \beta_0 \quad \text{and} \quad \alpha_1 > \beta_1
\]

Then based on (2), the present value of revenue from default \(-R\), which is expropriated contracted worker rents -- is given by the first integral in (2):

\[
R(t) = \int_0^t e^{-r(t-\tau)} d\tau + \int_0^T e^{-r(T-\tau)} d\tau
\]

\[
= \frac{(\alpha_0 - \beta_0)}{r} \left[ 1 - e^{-r(T-t)} \right] + \left( \alpha_1 - \beta_1 \right) \left[ \frac{t-T e^{-r(T-t)}}{r} + \frac{1 - e^{-r(T-t)}}{r^2} \right]
\]

Differentiating \( R \) with respect to \( t \), setting the result equal to zero and solving for the resulting revenue maximizing tenure level at which to default, yields:

\[
t^* = T + \frac{\ln \left( \alpha_1 - \beta_1 \right) - \ln \left[ r \left( \alpha_0 + \alpha_1 T - \beta_0 - \beta_1 T \right) + (\alpha_1 - \beta_1) \right]}{r}
\]

where \( t^* < T \) because \((\alpha_0 + \alpha_1 T - \beta_0 - \beta_1 T) = W(T) - V(T) > 0\).

We are primarily interested in the change in the optimal default tenure level with respect to \( r \), \( \partial t^*/\partial r \). The expression for this derivative is complicated and cannot be easily signed. However, the general effect of changes in \( r \) on \( t^* \) can be derived by evaluating \( t^* \) at \( r = 0 \) and as \( r \to \infty \). In the case of
zero discounting \((r=0)\), optimal defaulting in a revenue maximizing sense occurs at \((T-t')\). This follows from setting \(r=0\), and calculating the present value of the revenue stream:

\[ R(t) = (\alpha_0 - \beta_0) \int_t^T d\tau + (\alpha_1 - \beta_1) \int_t^T \tau \ d\tau = (\alpha_0 - \beta_0)(T-t) + \frac{(\alpha_1 - \beta_1)(T-t)^2}{2} \]

Solving for the optimal \(t\) yields \(t^* = T - \frac{(\beta_0 - \alpha_0)/{(\alpha_1 - \beta_1)} = T - t'}\). Note that if the optimal default point was decided upon at the beginning of employment, then the integral would be calculated from 0 to \(t\), yielding \(t^* = t'\). Similarly, it follows that as \(r \to \infty\) the revenue maximizing default point is at \(T\).

Specifically, evaluating the limit of (4) as \(r \to \infty\), clearly \(1/r\ln((\alpha_1 - \beta_1)\) approaches 0 as \(r \to \infty\), and using L'Hôpital's rule it follows that as \(r \to \infty\), then \(1/r\ln[r(\alpha_0 + \alpha_1 T - \beta_0 - \beta_1 T) + (\alpha_1 - \beta_1)] \to 0\), so that \(t^* \to T\) as \(r \to \infty\). Therefore, as \(r\) increases, the revenue maximizing point at which to default rises, so that in a period of economic disruption it is more likely that defaulting firms in need of revenue will default on workers with relatively longer tenure with the firm.\(^{10}\)

Firm defaults in a Lazear model may also proceed from a rise in the returns to default consequent on a technological shock, because a reduction in the value of workers' output will raise the costs to the firm of adhering to the contracted wage profile, thereby raising the returns from default. A technological

\(^{10}\)It may be reasonable to assume that default will only occur for workers currently receiving wages in excess of the value of their output -- i.e., workers for whom \(t > t'\). This allows us to considerably simplify (3) by only considering a revenue function that begins at \(t'\). We may then evaluate an expression with \(\alpha_0 = \beta_0\), which causes the first term in (3) to equal zero. Then revenue is increasing in \(T\) and decreasing in \(r\), namely:

\[
\frac{\partial R}{\partial T} = (\alpha_1 - \beta_1)(T-1) e^{-r(T-1)} > 0
\]

\[
\frac{\partial R}{\partial r} = (\alpha_1 - \beta_1)\left[ \frac{T e^{-r(T-1)} [1 + r (T-t)] - t}{r^2} + \frac{e^{-r(T-1)} [2 + r (T-t)] - 2}{r^3} \right] < 0
\]

In this case (4) may be written as:

\[
t^* = T - \frac{\ln (1 + r T)}{r} \Rightarrow \frac{\partial t^*}{\partial r} = \ln (1 + r T) - \frac{r T}{1 + r t} > 0
\]

which holds for all \(r > 0\).
shock that shifts the $V(t)$ function down by lowering $\beta_0$ increases the revenue from default:

$$\frac{\partial R}{\partial \beta_0} = -\frac{1}{r}[1 - e^{-r(T-t)}] < 0$$

This should unambiguously increase default profits and the probability of default, because it is not clear what mechanism would raise costs of default in this context. Differentiation of (4), however, suggests that a technological shock will lower the optimal tenure level at which to default:

$$\frac{\partial t^*}{\partial \beta_0} = \frac{1}{r (\alpha_0 + \alpha_1 T - \beta_0 - \beta_1 T)} > 0$$

$$\frac{\partial^2 t^*}{\partial r \partial \beta_0} = \frac{- (\alpha_0 + \alpha_1 T - \beta_0 - \beta_1 T)}{[r (\alpha_0 + \alpha_1 T - \beta_0 - \beta_1 T) + (\alpha_1 - \beta_1)]^2} < 0$$

The second partial, however, indicates that if $r$ is increasing due to the technological shock, this will generate a countervailing effect on $t^*$.

The effect of adverse economic conditions on the costs of default is less deterministic. Lazear (1981) assumes constant, one-period default costs. More generally, the exact shape of the cost of default function will depend on the discount rate used by workers and the information transmission process by which firm default is translated into reputation costs. For example, reputation costs might increase or decrease with the tenure level of the dismissed workers; high tenure workers generally lose fewer rents upon dismissal, but the outside labor force may perceive dismissal of high tenure workers as a more severe violation of the implicit agreement. Given this theoretical ambiguity concerning how costs change with tenure, we assume constant default costs ($\delta$) across all tenure levels and discount rates. To facilitate

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11 If workers have low discount rates then worker rents will tend to be higher near $t^*$. To the extent that workers can impose reputation costs on firms that are comparable to what the workers perceive are the value of their losses, costs will tend to be higher for the firm if they dismiss workers near mid-life with the firm. Alternatively, if workers highly discount the future (say, due to an adverse economic shock that leads them to be rather pessimistic about the survival likelihood of the firm), then costs will tend to rise with $t$ since the contemporaneous value of rents also tend to rise with $t$. 

9
our discussion of discounting and horizon effects on default, we extend Lazear's analysis by allowing for
default costs that may persist over many periods. The present value of costs for default at $t$ are thus:

$$C(t) = \delta \int_t^{T'} e^{-r(T'-t)} \, dt = \frac{\delta}{r} \left( 1 - e^{-r(T'-t)} \right) \quad (5)$$

where

$$\frac{\partial C}{\partial T'} = \delta \, e^{-r(T'-t)} > 0 \quad (5.1)$$

$$\frac{\partial C}{\partial r} = \frac{\delta}{r} \, e^{-r(T'-t)} \left[ \frac{1}{r} + (T'-t) \right] - \frac{\delta}{r^2} < 0 \quad (5.2)$$

From (5.1) we see that as $T'$ (time until firm failure) falls, default costs also fall. Similarly, from (5.2),
as the discount rate rises, default costs fall. The general implication is that adverse economic
conditions reduce default costs, thereby increasing default probabilities.

An additional cost of dismissal is a rehiring/re-training cost. Yet, as $r$ increases (or $T'$
decreases), this cost will approach zero, thereby unambiguously lowering costs. Similarly, if one cost
of cheating is the need to pay higher wages to attract new workers, then lower future hiring plans reduce
this cost, thereby leading to higher profits from default.

Overall, the delayed payments model suggests that sectoral decline increases default incentives
and raises the optimal tenure level at which to default. If sectoral decline is accompanied by
technological shocks, however, the tendency for firms to default on higher tenure workers will be
moderated.

(ii) Specific Human Capital

The implications of the specific human capital model for the effects of adverse economic
conditions on employee retention can be evaluated in a similar fashion. Assume that the interval $[0,t^*]$

\[\text{The inequality in (5.2) follows from } e^x > (1 + x) \text{ for } x > 0, \text{ where } x = r(T'-t).\]
represents the period over which firm-specific skills are acquired, and that \( W(t) \) and \( V(t) \) represent the worker's wage and productivity profiles in the current firm. Then:

\[
\begin{align*}
W(t) > V(t) & \text{ for } t < t'' \\
W(t) = V(t) & \text{ for } t = t'' \\
W(t) < V(t) & \text{ for } t > t''
\end{align*}
\]

The relative pattern in the wage and productivity profiles is the opposite of that in the delayed payments case. At higher tenure levels the value being produced by the worker exceeds wage payments, so that the firm will generally prefer to retain senior workers. The probability of worker dismissal is a decreasing function of the rents accruing to the firm, with no role for reputational costs due to premature dismissal.

For expository purposes, assume that after on-the-job training (OJT) is completed, \( V(t) \) jumps discontinuously above \( W(t) \). Assuming otherwise linear wage and productivity profiles implies:

\[
\begin{align*}
W(t) &= \omega_0 + \omega_1 t + \omega_2 D + \omega_3 t^* D \\
V(t) &= \nu_0 + \nu_1 t + \nu_2 D + \nu_3 t^* D \\
D &= 0 \text{ for } t < t'' \text{ and } D = 1 \text{ for } t \geq t''
\end{align*}
\]

All parameters are positive in the preceding expressions. Given sharing of the OJT costs and returns we know that \( \nu_0 < \omega_0 \); also, \( \nu_1, \omega_1, \nu_2, \) and \( \omega_2 \) have values such that \( V(t) < W(t) \) for \( t < t'' \), and \( V(t) > W(t) \) for \( t \geq t'' \). The relative slopes of \( V(t) \) and \( W(t) \) during the OJT period (\( \nu_1 \) and \( \omega_1 \)) are of little importance for our concerns. In the post-OJT period \( (t \geq t'') \) the relative slopes of \( V \) and \( W \), namely \( (\nu_1 + \nu_3) \) and \( (\omega_1 + \omega_3) \)

13Complete specific human capital contracts also specify the relationship between \( V(t) \) and alternative wages. For our purposes, however, we do not need to specify this relationship; it only influences separations through quits, which we do not analyze.

14Structuring the profiles in this fashion is not critical to the argument that follows, but it provides a useful framework for discussing the effects of different growth rates in productivity and discount rate changes on the rents accruing to firms from OJT investments. In the formal training case, a jump in \( V(t) \) at \( t'' \) may follow from the worker being given a new position and/or being given new types of physical capital to work with upon completion of the training program.
respectively, depend primarily on the degree of continuing enhancement of skills on the job (and also the
temporal pattern in the sharing of rents from the specific capital investments). In particular, \( \nu_1 + \nu_3 > \omega_1 + \omega_3 \) suggests relatively high continued skill enhancement, and \( \nu_1 + \nu_3 < \omega_1 + \omega_3 \) suggests relatively low continued skill enhancement.

As with the delayed payments model, we want to investigate the implications of this specific human capital model for the tenure pattern in dismissals under adverse economic conditions such as sectoral decline. These implications can be derived without resorting to formal comparative statics, for three distinct cases. The results are based on the change in rents over time, which in turn depend on the relative slopes of the \( V(t) \) and \( W(t) \) schedules during the post-OJT period \( t \geq t" \). First, if \( r = 0 \) (no discounting), the firm prefers to dismiss workers closest to retirement, regardless of the relative slopes of \( W \) and \( V \) (as long as \( V > W \)). Now consider the effect of an increasing discount rate \( (r) \); it is helpful to note that as \( r \rightarrow \infty \), only the current period matters. If \( \nu_1 + \nu_3 = \omega_1 + \omega_3 \), per-period rents to the firm are constant for all \( t \geq t" \), and increasing \( r \) makes employers increasingly indifferent concerning the tenure level at which to dismiss workers (for \( t \geq t" \)); relative to the case of \( r = 0 \), this should lower the average tenure level of dismissed workers. If \( \nu_1 + \nu_3 > \omega_1 + \omega_3 \), per-period rents to the firm are increasing for \( t \geq t" \), and as \( r \) increases the firm increasingly prefers to dismiss workers with lower tenure (for \( t \geq t" \)). If \( \nu_1 + \nu_3 < \omega_1 + \omega_3 \), per-period rents to the firm are decreasing for \( t \geq t" \), and as \( r \) increases the firm increasingly prefers to dismiss workers with higher tenure (for \( t \geq t" \)).

Thus, if enhancement of OJT skills is relatively ongoing \( (\nu_1 + \nu_3 \geq \omega_1 + \omega_3) \), adverse economic conditions -- in the form of an increased discount rate -- increase firms’ incentives for retention of senior versus junior workers. If enhancement of OJT skills is less ongoing, so that \( \nu_1 + \nu_3 < \omega_1 + \omega_3 \), adverse conditions will reduce firms’ incentives for retention of senior versus junior workers. Similar results hold if adverse conditions take the form of a decline in the firm’s anticipated operating horizon.

Adverse conditions may arise due to a technological or product demand shock that shifts down
the \( V(t) \) schedule. If the entire \( V(t) \) schedule shifts down -- through a decrease in \( \nu_0 \) -- then the incentives to retain senior as opposed to junior workers are unchanged. Alternatively, the shock may lower the value of existing firm-specific skills, which would reduce \( \nu_2 \) and perhaps \( \nu_3 \). By reducing rents produced by senior workers, such a shock will increase incentives for dismissal of senior workers.\(^{15}\)

Overall, the delayed payment and specific human capital models make somewhat different predictions concerning retention of senior employees under varying labor market conditions. Under delayed payments, retention of senior employees represents an equilibrium relationship based on the presence of net costs of default. Adverse economic conditions, as reflected by increased discount rates or a declining horizon, generally increase firm default incentives through reducing expected default costs. Also, adverse conditions tend to increase the optimal tenure level at which to default, although this effect may be moderated by technological shocks. Under specific human capital acquisition, retention of senior employees is generally preferred when skill enhancement is ongoing, but not if skill enhancement is relatively static or a technological shock lowers that value of existing firm-specific skills. The models are therefore only partially distinguished. Both predict weakening of the tenure effect on job security under adverse economic conditions when the value of senior workers' skills is constant or depreciating, possibly due to technological shocks. Therefore, to help distinguish between the models in the empirical work that follows, we incorporate data on industry technological change to control for any resulting effects on the value of firm-specific skills or the relative values of the productivity and contracted wage profiles.

Testing these propositions requires the use of firm separation data. Because employers rely heavily on temporary layoffs for employment adjustments (Feldstein 1975; Lilien 1980; Lilien and Hall 1986), we use the incidence of recall to the original employer in a sample of layoff unemployment spells.

\(^{15}\)For example, technological change may lead firms to retire older workers rather than invest in their retraining (see Mincer and Higuchi 1988; Bartel and Sicherman 1993a).
as our measure of worker/firm attachment. Under adverse economic conditions, the incentives to recall
senior workers should be reduced. Our primary intent is to examine whether such behavior is observable
in the data, and also to evaluate implications for the models discussed above, without necessarily
providing a strong test to distinguish between them.

We focused above on increases in the discount rate or decreases in firms' horizons as indicators
of adverse conditions. In the empirical work, we proxy for adverse conditions using industry/geographic
sector employment growth rates, under the assumption that declining sectors face adverse conditions of
the form we discuss. An alternative proxy for adverse conditions is the sectoral unemployment rate; this
provides a better measure of tightness in the relevant labor market than does employment growth, and
may therefore provide a useful proxy for the enforceability of delayed payment and specific human capital
contracts. We describe these data in detail in the next section, and the empirical tests in Section IV.

III. PSID Unemployment and Recall Data, Sector Employment and Technological Change Data

To test the implications of these models for the tenure pattern in recall of laid-off workers under
varying labor market conditions, we use a sample of unemployment spells derived from Waves XV and
XVI (1982-83) of the Panel Study of Income Dynamics (PSID). As described in Katz (1986), this data
set contains information on whether the respondent's most recent unemployment spell from the previous
year ended in a new job or in return to the original employer ("recall"). Our unemployment spell data
are derived from a sample of 2844 household heads (male and female) aged 22 to 65 in 1981 who were

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16 Our competing risk model of unemployment durations is estimated conditional on the occurrence of an
unemployment spell. Given the possibility of firm opportunism in both the recall and the initial layoff process,
however, these two events may not be independent. Estimating such a model is beyond the scope of this paper.

17 Katz (1986) used Waves XIV and XV (1981-82). We do not use the 1981 data for reasons described below
in the text. Data on recalls was not collected in the PSID after 1983. As discussed by Katz, unweighted use
of the Survey of Economic Opportunity low-income households does not appear to affect the unemployment duration
results.
labor force participants in the 1980-83 surveys; our unemployment data therefore exclude spells that end in retirement or other labor force withdrawal. The spells are matched with the respondent’s pre-spell personal and job characteristics, as described in detail below. Matching of unemployment spells with personal and job variables, and dropping spells that appear to be initiated by quits, yields a sample of 308 spells from the 1982 survey and 383 spells from the 1983 survey. Unlike previous authors, we stratify the sample by year in our empirical work. As discussed in the next section, the 1982 and 1983 results, which correspond to unemployment spells beginning in 1981 and 1982, differ substantially, probably due to the different points in the business cycle to which these years correspond.

A critical variable in our analysis is tenure in the firm. As discussed by other authors (see Brown and Light, 1992, for a comprehensive treatment), the PSID tenure measure is quite noisy. Much of this stems from definitional inconsistencies; prior to 1981, the distinction between tenure in the firm and tenure in a job was not clearly made. Since 1981, however, the distinction is relatively clear in the survey. Also, the tenure questions were not asked in 1980. We therefore do not use the Wave XIV (1981) data, for which pre-spell tenure corresponds to 1980. We lose some of the 1982 spell observations because pre-spell tenure is only available for respondents who were employed in the previous year, but this loss is small. Taking 1981 tenure as correct, we use the job change information in the PSID to construct a longitudinally consistent pre-spell tenure measure (tenure in 1982) for the 1983 data.

It is also critical in our study to use appropriate measures of labor market conditions. As argued in the preceding section, firms’ incentives for retention of senior workers may be reduced in industries that are experiencing long-run decline or are cyclically sensitive. To measure labor market conditions,

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18 We lose 62 spells from our 1982 sample and 43 spells from our 1983 sample because these individuals identified quitting as the reason for the corresponding job change between surveys. We delete these observations because our focus is on firms’ incentive for employee retention. Of the quitters, 20-25% are also identified as recalls, which suggests that employer retention incentives may be relevant for these observations. Our main results, however, are not sensitive to the exclusion of quits from the sample. Results derived from a sample that includes quits are available on request.
we tabulated several variables from the March Current Population Survey (CPS) Annual Demographic Files for various years and matched them to our PSID data. The CPS files provide current information on employment and extensive information on labor market experience over the previous year. To measure the extent of sectoral decline, we use the change in sector employment levels for the 5- and 10-year periods preceding our sample frame. The sectors are defined by respondents' reported industry of longest job in the previous year and geographic location of current residence; we use 43 detailed industry categories and 9 geographic division categories, which produces 387 sectors. As mentioned in the preceding section, we are also interested in the effects of sector unemployment on employee retention. We therefore use the industry/geographic sector unemployment rate, calculated as total employment weeks divided by total labor force weeks in the sector, based on a sample of 18 to 64 year-olds. Use of these sectoral variables provides an additional reason why we do not use the 1981 PSID spell data; the corresponding pre-spell industry variables (from 1980) are 2-digit and do not enable an exact match with our CPS data.

An additional data issue is the possibility of incorrect assignment of pre-spell values. The unemployment spells refer to the respondent's most recent spell originating in the year prior to the survey year. For the personal characteristic variables, such as age and education, mismatches will be minimal. Our pre-spell industry and tenure measures, however, which are the focus of our analysis, are less precise than the personal characteristic variables. A primary reason for this is the possibility of multiple job-changing between survey dates. Our assignation procedure for the pre-spell industry and tenure measures

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19 Murphy and Topel (1987) use a similar measure -- in the aggregate -- to analyze U.S. unemployment during 1968-85. We do not use the PSID county unemployment rate variable because our preliminary investigations suggested that it is not an appropriate measure of slackness in the relevant labor market.

20 All CPS tabulations were weighted by the March supplement weights. We treat the tabulated variables as fixed in our sample. To mitigate potentially excessive sampling error and mismatches due to geographic mobility between the survey date and the longest job in the previous year, we also tabulated the unemployment and employment change variables by industry sectors only; the results from using the industry sector variables are similar to those using the industry/geographic sector variables (and are available on request).
assumes that the spell originated from the respondent’s job in the previous survey year (or most recent job, for respondents who were unemployed in the previous survey year). For individuals with multiple spells between survey dates, this assumption is likely to be incorrect. The resulting bias in the estimated coefficients on the tenure and industry/division labor market variables, however, should correspond to the classical measurement error case; the coefficients will be biased toward zero, and our results will understate the true size and significance of the effects.  

Finally, as described in Section II, our results may be sensitive to differential rates of technological change across industrial sectors. We therefore use technological change variables similar to those used by Bartel and Sicherman (1993a, 1993b), based on estimates provided by Jorgenson et al. (1987). Jorgenson et al. estimated yearly productivity growth for 35 U.S. industry sectors over the period 1947-85. Although our 43 industry categories are somewhat more detailed than Jorgenson et al.’s, and our sectors are further disaggregated by 9 geographic regions, the Jorgenson et al. data provide a reasonable industry match and were the best available measures of technological change for workers’ pre-spell industry in our sample. To measure the rate of long-run technological change, we formed variables that represent the 5- and 10-year average productivity growth rate by industrial sector. Bartel and Sicherman also used variables that capture the unexpected change in the rate of technological change. Like them, we measure deviations from trend productivity growth by forming variables that represent the current year deviation from 5- and 10-year productivity growth, normalized in the form of a Z-score.  

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21 Preliminary estimation with the sample restricted to individuals who experienced a single spell of unemployment between surveys did not appreciably change the results.

22 We thank Ann Bartel and Nachum Sicherman for providing machine-readable data files. Like Bartel and Sicherman (1993a, 1993b), we assume that productivity growth rates adequately proxy for rates of technological change.

23 Bartel and Sicherman only use 10-year productivity growth variables; we also use 5-year growth. The Z-score is the current period’s deviation from mean productivity growth divided by the standard deviation of productivity growth, both calculated over the same period.
In section V, we refer to these latter variables as technological "shock" variables.

Table 1 presents tabulations that illustrate the relationships between recall incidence, tenure, and industry/geographic sector employment growth in our PSID/CPS data. Termination of unemployment spells through recall is frequent; 52% of the spells end in recall (53.1% in the 1983 sample and 50.3% in the 1982 sample). New job finders comprise 25.6% of the 1983 sample and 34.1% of the 1982 sample. The remaining observations (21.3% in 1983 and 15.6% in 1982) are censored. The recall rate is higher and the new job acceptance rate is lower in our sample than in Katz’s, due to our more restrictive sample selection criteria. The table tabulates percentage recall by pre-spell tenure categories and high/low sectoral employment change over the 10 years preceding each sample period. The sample shares indicate that over a third of the sample in each year have at least 5 years of pre-spell tenure, and 16-20% of each sample have at least 12 years of pre-spell tenure. A strong positive relationship between tenure and recall incidence appears in both samples, in high-growth and low-growth sectors. Within tenure categories, recall incidence is typically higher in low-growth industries in the 1983 sample, but varies across high-growth and low-growth industries in the 1982 sample. No particular pattern emerges in regard to the strength of the tenure effect on recall across high-growth and low-growth sectors. Although these results are suggestive of potential links between recall incidence, tenure, and sectoral employment growth, drawing more reliable conclusions requires explicitly accounting for new job finding and censoring in a multivariate framework, using the competing risks model described in Section IV.

IV. Competing Risks Specification

Use of a competing risks framework allows us to more carefully analyze the incidence and pattern

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24 Some of these workers may have chosen layoff. As noted by Medoff (1979), many union contracts allow senior union workers to waive their job rights and be laid-off while retaining their seniority rights to rapid recall. Implicit agreements with nonunion workers may produce similar behavior.
of recalls and new job finding in our sample of unemployment spells. Katz (1986) presents a model of
d job search with positive recall probabilities that leads to a competing risks specification of unemployment
spell durations. This specification distinguishes between two routes of exit from unemployment: recall
to the original employer, and new job acceptance. The competing risks specification is required because
the processes underlying these two events are likely to differ, implying different covariate effects and
different baseline hazards for new job and recall durations; these separate effects are confounded in single
risk analyses of unemployment spells. Existing empirical work confirms these hypotheses.25

Our primary focus is on pre-spell tenure and industry labor market conditions, and how they
affect employers’ recall probabilities. These factors are likely to affect recall and new job probabilities
differently; recall depends primarily on the incentives faced by the employee’s current firm, whereas new
job finding depends on employee and other firm behavior. The competing risk specification is therefore
ideal for our analysis. Following Katz (1986), let D_r and D_n be independent random variables (across
individuals) that represent the latent failure time for exit from unemployment through recall or new job
acceptance. Then if D_u = actual spell duration,

$$D_u = \min (D_r, D_n).$$

We assume a proportional hazards specification in continuous time for covariate effects; although
the unemployment durations are grouped into discrete intervals (weeks), the continuous time specification
provides a reasonable approximation. The cause-specific hazard functions for recalls and new job
acceptance are:

$$h_r(d, X) = \Pr( D_r = d | D_r \geq d ) = f(d)/(1-F(d)) = h_{0r}(d)e^{X_r}$$

$$h_n(d, X) = \Pr( D_n = d | D_n \geq d ) = g(d)/(1-G(d)) = h_{0n}(d)e^{X_n}$$

25 See Katz (1986), Han and Hausman (1990), and Sueyoshi (1991) for estimation of competing risks models
using PSID unemployment data. Katz and Meyer (1990) perform additional analyses incorporating recall
expectations in a sample of UI recipients from Missouri and Pennsylvania, as does Anderson (1992) using UI data
from New Jersey. Kalbfleisch and Prentice (1980, chapter 7) provide a detailed discussion of the competing risks
framework.
where \( d \) is an arbitrary spell duration, \( f \) and \( g \) are the probability density functions associated with \( D_r \) and \( D_n \), \( F \) and \( G \) are the corresponding cumulative distribution functions, \( X \) is a vector of covariates, and \( \beta \) and \( \gamma \) are corresponding coefficient vectors. The covariates operate through shifting the baseline hazard functions \( h_{d0}(d) \) and \( h_{d0}(d) \), which are specified below.\(^26\)

The corresponding survivor function for \( D_n \) -- \( \text{Pr}(D_n \geq d) \) -- is the product of the integrated hazards for recall and new job acceptance. The resulting log-likelihood function for the complete sample of recall durations, new job durations, and censored spells contains separate terms for the recall and new job distributions; these terms depend only on the parameters of the corresponding hazard. Estimation of the recall hazard parameters then proceeds as a standard single risk estimation, treating new job durations as censored; a similar single risk model, which treats recall observations as censored, is used to estimate the new job hazard parameters (see Kalbfleisch and Prentice 1980 or Katz 1986 for details). Our estimates are obtained through standard maximization techniques applied to the two resulting log-likelihood functions.

A potential drawback to this specification is its imposition of independent risks -- i.e., the purely stochastic components of the new job and recall hazards are assumed to be independent; for example, conditional on the covariates, early recall for an individual implies nothing about that individual's new job prospects. This restriction may not hold. Han and Hausman (1990) test the independence assumption, however, and find that it cannot be rejected in Katz's original data. Fallick (1992), using 1984 and 1986 Displaced Worker Survey data, also fails to reject independence in his competing risks analysis of nonemployment spells of industry switchers and industry stayers. We therefore do not investigate this issue further here.

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\(^{26}\) Sueyoshi (1991) statistically rejects the proportional hazards specification against time-varying coefficients in the Katz data. The non-proportionality applies only to the UI-receipt variable, however, probably because UI receipt is partially conditional on spell length (some individuals with very short spells do not have time to apply for and receive UI).
The baseline hazards remain to be specified. In the estimation, we use the Weibull specification:

\[ h_0(d) = a d^{a-1} \]
\[ h_\infty(d) = b d^{b-1} \]

This specification imposes a monotonic baseline hazard over the spell; the baseline is increasing or decreasing depending on whether the estimated duration dependence parameters \( a \) and \( b \) are greater or less than one. Because Han and Hausman (1990) formally test and reject monotonicity in the Katz data, we also estimated our models using the partial likelihood approach of Cox (1972), which imposes no restrictions on the shape of the baseline hazard function. This model produced results that are similar (in direction and significance) to those reported for the Weibull specification (and are available on request). We therefore focus on the Weibull results, which are easier to interpret. The resulting duration dependence parameters may be biased downward due to unobserved heterogeneity in the data. Because we focus on specific parameter estimates rather than duration dependence, we do not model unobserved heterogeneity.\(^{27}\)

V. Empirical Results

We first focus on competing risk regression models that employ our measures of industry/geographic sector employment growth. Table 2, which uses the 1983 data, presents means for the independent variables in the first column, and regression parameter estimates for the new job and recall hazards in the second and third columns. The 1982 results are in Table 3. Due to our sample restrictions, the samples are somewhat older, have higher average tenure, and have fewer women than would representative labor market samples.

\(^{27}\) As discussed in Kiefer (1988), parameter estimates may be affected by both the heterogeneity distribution and baseline hazard function specified, although he argues that the latter is likely to be more important. Han and Hausman (1990) present a method for competing risks estimation that combines a parametric heterogeneity distribution with a non-parametric baseline hazard, but they find that the addition of heterogeneity has very little effect on the results.
The regression coefficients indicate the variables' effects on the hazard function; a positive coefficient increases the probability of exit, thereby decreasing spell duration. The results for the control variables are reasonably consistent with those from previous studies and are therefore only discussed briefly here. Education increases the new job hazard but may reduce recall prospects. Conditional on tenure, age does not affect new job finding or recall. Marriage increases the recall hazard in the 1982 data. Blacks have substantially lower new job hazards in 1983 but similar recall prospects to whites. UI receipt reduces the new job hazard and also the 1982 sample recall hazard.\textsuperscript{28} Union coverage substantially increases the recall hazard, but substantially reduces the 1983 new job hazard.\textsuperscript{29} Average firm size in workers' pre-spell 2-digit industry (tabulated from the 1979 May CPS) does not significantly affect the recall or new job hazards in either year.\textsuperscript{30} The duration dependence parameters indicate a declining recall hazard (particularly in 1983), but the hypothesis of a flat new job hazard is not rejected.

Our focus is on the estimated tenure and sectoral employment change variables, along with their interaction. Pre-spell tenure has significant positive effects on the 1983 and 1982 recall hazards and a significant negative effect on the 1982 new job hazard. These results are consistent with either a specific

\textsuperscript{28}Stratification of the full samples into UI-receipt and non-UI sub-samples does not change our substantive results; the results are primarily driven, however, by the UI recipient sample.

\textsuperscript{29}This is consistent with Medoff's (1979, p. 387) finding that "...the greater use of layoffs in unionized establishments during periods of decreased commodity demand permits a higher rate of rehires and lower rate of new hires when commodity demand increases." Stratification of the sample into union and non-union sub-samples indicated that our main results are substantially stronger in the non-union sample (results available on request). In particular, the reduction in the tenure effect on recall due to adverse conditions (discussed below) does not occur in the union sample, presumably because unions have explicit mechanisms to protect the job security of senior workers.

\textsuperscript{30}The firm size effect on recall is positive and significant when the sector unemployment rate is used in place of sector employment growth (Table 5), and in all models when we use firm size tabulated at the 3-digit rather than the 2-digit industry level. None of the other coefficients, however, are sensitive to the choice of firm size variable. Because the 3-digit measure causes the loss of 3 observations, we use the 2-digit measure; results from models that use the 3-digit measure are available on request. Also, although larger firms may be more likely to use delayed payment contracts due to stronger reputational effects (Parsons 1986), they also tend to provide higher training investments (Barron et al. 1989; Brown and Medoff 1989; Idson 1989). Stratification of the sample by average firm size, however, indicated that our main results do not vary by firm size.
human capital or delayed payment contract explanation. First, the shared investment or implicit agreement provides strong recall incentives. This is particularly true under specific human capital accumulation, but should also be true in the delayed payment case: a firm that uses layoffs as an excuse to permanently discharge workers with high tenure (at the same or greater rate as workers with low tenure) is likely to suffer reputational costs associated with defaulting on delayed payment contracts. Also, the excess of wages over alternative marginal product for high tenure workers, which is predicted by both theories, is likely to reduce new job acceptance (see Valletta 1991). It is not immediately clear why the tenure effect on the new job hazard differs in the 1982 and 1983 data. However, in models that replace sectoral employment growth by the sectoral unemployment rate (see Table 5), tenure has a significant negative effect on new job finding in the 1983 data.

The 1982 and 1983 sample results differ significantly in regard to the effects of sector employment growth and the tenure/employment growth interaction. We focus first on the 1983 results, and return to the 1982 results below. The sectoral employment growth variable (the difference in log sector employment between the two periods, which can be negative) has a strong negative effect on the recall hazard but a strong positive effect on the new job hazard in the 1983 data. The positive effect on new jobs suggests that workers laid-off in industries experiencing rapid employment growth quickly find re-employment in the same industry; this interpretation is consistent with Fallick's (1992) analysis of industry switchers and stayers. In contrast, higher turnover in rapidly expanding industries may reduce recall probabilities for workers laid-off in such industries, thus producing the negative effect on the recall hazard.

The key result is the significant positive coefficient on the interaction between pre-spell tenure and sector employment growth in the 1983 recall equation. This positive interaction effect suggests that firms in declining sectors are less likely to recall high tenure workers than are firms in growing sectors. As argued in Section II, this result is consistent with employer default on delayed payment contracts,
because sectoral decline is likely to increase the net benefits of such defaults. The result may also be consistent, however, with a specific human capital model that incorporates continued skill enhancement and allows for possible effects of technological change.

Although the tenure effects on recall are similar in the 1982 and 1983 data, the employment growth and tenure/sector employment growth interaction effects appear only in the 1983 data. Neither coefficient even approaches reasonable significance in Table 3 (1982 sample). Because the spells occurred in the year prior to the survey, and because the overall unemployment rate was substantially higher in 1982 than in 1981, the results suggest that the effects of industry conditions on the tenure pattern in employee retention incentives depend largely on overall labor market conditions. These results also suggest that the sector unemployment rate may be an additional important measure of sector labor market conditions; we investigate this below, in Table 5.

To investigate the possibility that our 1983 sample results are caused by technological changes that lower the value of firm-specific investments, we introduced the industry technological change and shock variables, as described in Section III, into our competing risk equations. We controlled for the 5- and 10-year mean productivity growth rates in the worker's industry sector, and the deviation from mean growth (the technological "shock") expressed in the form of a Z-score. We also included interactions between the technological change variables and tenure. If the reduction in the tenure effect on recall is caused by technological changes or shocks in workers' pre-spell industrial sector, then interactions between tenure and the technological variables should reflect this, and their inclusion should reduce or eliminate the significant interaction between tenure and sector employment growth.

We estimated the models with the various technological change and shock variables included separately, then together, along with their tenure interactions. None of these specifications altered the

31 Medoff and Abraham's (1981) evidence from one large company is consistent with this interpretation. They found that although seniority had no effect on involuntary terminations under normal conditions, during high layoff periods the probability of involuntary termination is positively related to seniority.
key conclusions based on Table 2. In particular, the tenure/technological change (or shock) interactions never approached significance, and the coefficient on the interaction between tenure and sector employment growth was not affected by inclusion of the technological variables and their interactions with tenure. For illustrative purposes, Table 4 presents results based on inclusion of the 5-year shock variable, which was the only technological variable that produced a significant coefficient in the models. Its coefficient reveals a positive effect of the current period's deviation from 5-year productivity change on recall probabilities. Compared to the results in Table 2, the negative coefficient on sector employment growth in the recall equation is reduced somewhat in size and significance due to inclusion of the technological variable. These two results probably reflect a negative correlation between sector employment growth and technological shocks. Most importantly, the tenure/sector employment growth interaction coefficient remains essentially unchanged in size and significance relative to Table 2. Although our technological change variables are rough at best, the robustness of our main results with respect to their inclusion supports the view that the 1983 results are consistent with firm defaults on delayed payment contracts.

Discussion of the magnitudes implied by the Table 2 results also supports this interpretation. In particular, the tenure coefficients imply large effects on expected median recall duration (calculated at the mean values for all covariates), but the tenure effect on recall is substantially mitigated by sectoral employment decline. In the 1983 data, each additional year of tenure reduces expected recall duration by 1.0 weeks; the corresponding median duration is 22.3 weeks, which implies that each additional year of tenure has approximately a 4.5% effect on median recall duration. Each additional ten percentage points of 10-year sectoral employment decline reduces 1983 recall duration by 2.3 weeks, but also reduces the positive tenure effect on recall by .011. The standard deviation of the 10-year employment growth rate in the 1983 sample is .306. Thus, a one standard deviation decrease in the sector employment growth rate reduces the tenure coefficient (.047) in the recall equation by approximately
0.033. These estimates imply that the total tenure effect on the recall hazard is negative in sectors with
greater than 35% employment decline over 1972-1982, which is within the range of the 10-year sector
employment growth variable in our 1983 sample (approximately at the 7th percentile of the distribution
of this variable). This sign reversal for the tenure effect on recall incidence in declining sectors seems
inconsistent with a specific human capital interpretation, which reinforces our interpretation of the results.

Table 5 presents results for the 1983 sample with the 5-year sector employment growth rate and
the 1982 sector unemployment rate replacing the 10-year employment growth rate as measures of sectoral
labor market conditions. Results for the other covariates are very similar to those from Table 2 and
therefore are not reported; also, the substantive results for 1982 are similar to those in Table 3 and
therefore are not reported.

The sector unemployment rate reduces the new job hazard, although the effect is not as precisely
estimated as one might expect; it only attains significance at the 13% level in the 1983 data (but at the
2% level in the 1982 data). Although this result appears problematic, Fallick’s (1992) results from the
1984 and 1986 Displaced Worker Surveys suggest that nonemployed workers attached to slack industries
increase their search intensity and reduce their reservation wages in other industries; this behavior may
mitigate the expected negative effect of industry unemployment on new job finding in our data. In
contrast, sector unemployment has a strong positive effect on recall prospects in the 1983 data,
presumably due to a high incidence of short, explicitly temporary layoffs in slack sectors during a
recession. The positive effect of sector unemployment on recall prospects, however, does not appear

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32 We also tried an industry failure rate variable (as in Curme and Kahn 1990; obtained from Dun and
Bradstreet’s Business Failure Record) as an additional proxy for labor market conditions. Although its sign was
mostly as expected in the estimated equations, the coefficients were imprecisely estimated, and therefore are not
reported (results available on request).

33 Lilien (1980, p. 28) finds “… considerable cyclical variation in the rehire rate, with 78% of the layoffs in
the recessionary year of 1975 generating rehires, as compared with 60% in 1970.” Also, Lilien and Hall (1986,
pp. 1009-10) note that “… a continual process of recurrent layoffs is one of the ways that work sharing operates
during a period of sustained slack … both layoffs and rehire continue at high rates after the economy has reached
in the 1982 data (results available on request); this is similar to the 1982 sample results reported above (Table 3), where the sector employment growth rate does not affect recall prospects. Finally, note that in Panel B of Table 5, the tenure effects on recall and new job finding are substantially strengthened in size and significance relative to those in Table 2. This suggests that sectoral unemployment is less closely correlated with individual worker tenure than is sectoral employment growth.

As in Table 2, our primary focus is on the tenure interaction terms in the recall equations. This interaction is positive for the 5-year growth variable, but is smaller and less reliably estimated (significant at the 9% level) than for the 10-year change in Table 2. The stronger results for the 10-year industry growth rate suggests that a 10-year horizon may be more relevant for the default considerations discussed in Section II. In the bottom panel, the significant negative coefficient on the interaction between tenure and sector unemployment has the same interpretation as the positive coefficient on the tenure/employment growth interaction variable from Table 2; higher unemployment implies a slacker labor market, as does lower employment growth.\(^{34}\) Thus, the unemployment rate and tenure interaction results suggest that firms operating in high unemployment sectors are less likely to recall high tenure workers than are firms in low unemployment sectors. As with the results from Table 2, our Table 5 results are invariant to the inclusion of technological change variables. In addition, tabulations based on the tenure and tenure/unemployment interaction coefficients suggest that the tenure effect on recall is zero in sectors with approximately 19% unemployment, which is within the range of the sector unemployment variable in our 1983 sample (approximately at the 95th percentile of the distribution of this variable). Because the unemployment rate may raise net default benefits through lowering reputational costs, this result appears

\(^{34}\) Although the employment growth variables partially reflect business cycle effects, their pure trend growth component is higher for the 10-year change, for which we obtain stronger results. Our employment growth results are therefore not driven entirely by variation in unemployment rates, and vice versa.
VI. Conclusions and Directions for Future Work

Our competing risks analysis of unemployment spells indicates that tenure effects on employee retention are significantly weakened by adverse economic conditions. In particular, an estimated strong positive tenure effect on the probability of recall from layoff is significantly mitigated by labor market slackness, where the latter is measured by relative sectoral (i.e., industry/geographic division) employment growth and unemployment. We interpret this evidence as being consistent with a delayed payments model in which sectoral slackness increases the net benefits of permanently dismissing -- or, in our analysis, not recalling -- senior workers. Although this result could also be consistent with a specific human capital view in which sectoral slackness arises from technological changes that reduce the value of existing firm- and industry-specific skills, the invariance of our results with respect to the incorporation of data on industry technological change suggests otherwise. As noted, however, our intent was not to provide a precise test between these views. Future investigation of this hypothesis using data that allows for more precise empirical tests would be useful.

Various avenues for related work are suggested by this study. Use of direct measures of on-the-job training (available in both the PSID and NLS) in our analysis of recall decisions may generate more precise results with regard to the role specific human capital in firm separation decisions. Researchers might also examine differential effects across occupations, under the assumption of differential on-the-job training levels and monitoring difficulty across occupations. Given that firm opportunism may extend

\[\text{35Our tenure interaction coefficients might also reflect the influence of widespread plant closures in declining sectors. Since plant closures should eliminate any recall probability for workers at all tenure levels -- except in rare cases where seniority rights carry across plants in multiprant firms -- they may dampen the positive effect of tenure on recall probabilities, thereby producing our tenure interaction coefficients. Closure of plants with high average worker tenure would reinforce this effect, but such a pattern of closures might itself reflect firm default on implicit contracts. Unemployment spells arising from plant closures only comprise about 8% of our sample, however, and our substantive results are mostly invariant to their removal from the sample (results available on request).}\]
beyond the recall decision to layoff and separation behavior in general, extension of our research to
general firm separation decisions may prove fruitful.

Expansion of our time frame deeper into the 1980s may also yield interesting results. Our results
suggest the presence of firm opportunism in regard to layoff unemployment spells beginning in 1982, the
first year of a severe recession. Because the opportunism is in regard to senior workers, our results
provide some empirical verification of the common claim that older workers -- particularly those who
stayed at one firm for a long time -- faced severe employment problems during the 1980s. Additional
data may enable tests of whether such behavior continued into 1983, when the recession hit its low point.
It is also interesting to ask whether employer opportunism might have continued after the economy
recovered. Industrial demand shifts, takeovers and mergers, and firm downsizing continued to be
important issues during the boom of the 1980s, and contributed to a changing labor market. Under these
conditions, the incentive for opportunistic behavior with regard to retention of senior workers may have
continued throughout the decade. For example, it would be interesting to analyze the tenure patterns in
discharges or re-hires when a firm is acquired by another firm,\textsuperscript{36} because the acquiring firm is ostensibly
not bound by the implicit contracts of the former owner (Schleifer and Summers 1988).\textsuperscript{37}

The exact welfare implications of employer opportunism are unclear at this point and might be
a fruitful area for further investigation. Cornwell et al. (1991) find limited evidence for employer
opportunistic behavior through discharging workers covered by pensions during the 1980s. Such
opportunism may occur more generally, however, particularly under adverse conditions such as those
examined here. At a minimum, worker retention in a changing labor market is an important issue for
further study.

\textsuperscript{36} See Santiago (1987) for an analysis of this process for a particular acquisition. Because he examines a
unionized setting, however, the ability of employers to act either opportunistically or in a fashion consistent with
strict profit maximization is reduced.

\textsuperscript{37} Employment contract breach during an acquisition may reduce productivity. In the delayed payment case,
contract breach will reduce workers' willingness to accept backloading of pay, which will increase shirking; in the
specific human capital case, workers will be less willing to share in on-the-job training costs.
BIBLIOGRAPHY


Table 1 -- Percentage Recall,¹ by Pre-Spell Tenure and Industry/Geographic Sector 10-Year Employment Growth Rate

1983 Sample (N=382) -- Percentage Recall (.531 for full sample)

<table>
<thead>
<tr>
<th>Pre-Spell Tenure (share of sample)²</th>
<th>Δ Sector Employment &gt; average³</th>
<th>Δ Sector Employment &lt; average</th>
</tr>
</thead>
<tbody>
<tr>
<td>&lt; ½ year (.147)</td>
<td>.235</td>
<td>.136</td>
</tr>
<tr>
<td>½-1 years (.133)</td>
<td>.320</td>
<td>.577</td>
</tr>
<tr>
<td>1-3 years (.167)</td>
<td>.372</td>
<td>.571</td>
</tr>
<tr>
<td>3-5 years (.167)</td>
<td>.400</td>
<td>.529</td>
</tr>
<tr>
<td>5-12 years (.183)</td>
<td>.593</td>
<td>.744</td>
</tr>
<tr>
<td>12+ years (.202)</td>
<td>.800</td>
<td>.822</td>
</tr>
</tbody>
</table>

1982 Sample (N=308) -- Percentage Recall (.503 for full sample)

<table>
<thead>
<tr>
<th>Pre-Spell Tenure (share of sample)</th>
<th>Δ Sector Employment &gt; average</th>
<th>Δ Sector Employment &lt; average</th>
</tr>
</thead>
<tbody>
<tr>
<td>&lt; ½ year (.185)</td>
<td>.158</td>
<td>.105</td>
</tr>
<tr>
<td>½-1 years (.088)</td>
<td>.105</td>
<td>.375</td>
</tr>
<tr>
<td>1-3 years (.221)</td>
<td>.486</td>
<td>.451</td>
</tr>
<tr>
<td>3-5 years (.159)</td>
<td>.739</td>
<td>.500</td>
</tr>
<tr>
<td>5-12 years (.185)</td>
<td>.545</td>
<td>.743</td>
</tr>
<tr>
<td>12+ years (.162)</td>
<td>.833</td>
<td>.844</td>
</tr>
</tbody>
</table>

¹ New job finders and censored observations comprise the remainder for each category.

² The tenure categories are non-inclusive of the upper limit. Sample share in parentheses.

³ Mean of the industry/geographic sector employment growth rate (from 1972 to 1982 or 1971 to 1981) across all individuals in the sample; its value is .078 in 1982 and .185 in 1981.
Table 2 -- Competing Risks Model for Recall and New Job Hazards, 1983 Data, with 1972-1982 Sector Employment Growth Variables (variable means and coefficients; standard errors in parentheses)

<table>
<thead>
<tr>
<th>Variable</th>
<th>Mean</th>
<th>Recall Hazard</th>
<th>New Job Hazard</th>
</tr>
</thead>
<tbody>
<tr>
<td>Education</td>
<td>11.3</td>
<td>-.099</td>
<td>.155*</td>
</tr>
<tr>
<td>Age</td>
<td>36.7</td>
<td>.009</td>
<td>-.007</td>
</tr>
<tr>
<td>Married</td>
<td>.686</td>
<td>.458</td>
<td>.174</td>
</tr>
<tr>
<td>Race</td>
<td>.369</td>
<td>-.056</td>
<td>-.695*</td>
</tr>
<tr>
<td># of children (age &lt; 18)</td>
<td>1.30</td>
<td>.155</td>
<td>.082</td>
</tr>
<tr>
<td>Female</td>
<td>.141</td>
<td>.754</td>
<td>.058</td>
</tr>
<tr>
<td>UI Receipt</td>
<td>.725</td>
<td>-.380</td>
<td>-.418*</td>
</tr>
<tr>
<td>Union contract</td>
<td>.458</td>
<td>.829*</td>
<td>-.540*</td>
</tr>
<tr>
<td>Average firm size (2-digit industry)</td>
<td>1059.6</td>
<td>.00049</td>
<td>-.00049</td>
</tr>
<tr>
<td>Tenure at firm</td>
<td>6.23</td>
<td>.047*</td>
<td>-.024</td>
</tr>
<tr>
<td>ΔLog(Industry/Geographic Sector Employment)</td>
<td>.077</td>
<td>-.984*</td>
<td>1.17*</td>
</tr>
<tr>
<td>Tenure * ΔLog(Sector Emp.)</td>
<td>-.168</td>
<td>.106*</td>
<td>-.0110</td>
</tr>
<tr>
<td>Constant</td>
<td>1</td>
<td>-4.32*</td>
<td>-4.70*</td>
</tr>
</tbody>
</table>

(continued)
Table 2 (continued)

<table>
<thead>
<tr>
<th>Duration Dependence</th>
<th>0.657*</th>
<th>1.09*</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(.062)</td>
<td>(.101)</td>
</tr>
</tbody>
</table>

| Log-L               | -541.1 | -244.1 |

| Observations        | 382    | 382    | 382    |

Note: Independent variable values are pre-spell, hence from 1982.

* - denotes significance at the .05 level, two-tailed test.
Table 3 -- Competing Risks Model for Recall and New Job Hazards, 1982 Data with 1971-1981 Sector Employment Growth Variables (variable means and coefficients; standard errors in parentheses)

<table>
<thead>
<tr>
<th>Variable</th>
<th>Mean</th>
<th>Recall Hazard</th>
<th>New Job Hazard</th>
</tr>
</thead>
<tbody>
<tr>
<td>Education</td>
<td>11.4</td>
<td>-.094*</td>
<td>.114*</td>
</tr>
<tr>
<td>Age</td>
<td>34.5</td>
<td>-.015</td>
<td>-.014</td>
</tr>
<tr>
<td>Married</td>
<td>.623</td>
<td>.748*</td>
<td>.377</td>
</tr>
<tr>
<td>Race (Black = 1)</td>
<td>.432</td>
<td>.126</td>
<td>-.332</td>
</tr>
<tr>
<td># of children (age &lt; 18)</td>
<td>1.30</td>
<td>-.051</td>
<td>-.010</td>
</tr>
<tr>
<td>Female</td>
<td>.140</td>
<td>.951*</td>
<td>.312</td>
</tr>
<tr>
<td>UI Receipt</td>
<td>.701</td>
<td>-.725*</td>
<td>-.424*</td>
</tr>
<tr>
<td>Union contract</td>
<td>.464</td>
<td>.999*</td>
<td>-.172</td>
</tr>
<tr>
<td>Average firm size (2-digit industry)</td>
<td>1040.5</td>
<td>.00003-.00004</td>
<td>(.0003)</td>
</tr>
<tr>
<td>Tenure at firm</td>
<td>4.32</td>
<td>.093*</td>
<td>-.111*</td>
</tr>
<tr>
<td>ΔLog(Industry/Geographic Sector Employment)</td>
<td>.185</td>
<td>.060</td>
<td>.751</td>
</tr>
<tr>
<td>Tenure * ΔLog(Sector Emp.)</td>
<td>.604</td>
<td>-.025</td>
<td>.063</td>
</tr>
<tr>
<td>Constant</td>
<td>1</td>
<td>-2.92*</td>
<td>-4.21*</td>
</tr>
</tbody>
</table>

(continued)
Table 3 (continued)

<table>
<thead>
<tr>
<th>Duration Dependence</th>
<th>0.796*</th>
<th>1.15*</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(.080)</td>
<td>(.104)</td>
</tr>
<tr>
<td>Log-L</td>
<td>-386.8</td>
<td>-222.5</td>
</tr>
<tr>
<td>Observations</td>
<td>308</td>
<td>308</td>
</tr>
</tbody>
</table>

Note: Independent variable values are pre-spell, hence from 1981.

* - denotes significance at the .05 level, two-tailed test.
Table 4 -- Competing Risks Model for Recall and New Job Hazards, 1983 Data, with Technological Shock Variables (variable coefficients; standard errors in parentheses)

<table>
<thead>
<tr>
<th>Variable</th>
<th>Recall Hazard</th>
<th>New Job Hazard</th>
</tr>
</thead>
<tbody>
<tr>
<td>Tenure at firm</td>
<td>.047*</td>
<td>-.021</td>
</tr>
<tr>
<td></td>
<td>(.013)</td>
<td>(.022)</td>
</tr>
<tr>
<td>ΔLog(Industry/Geographic Sector Employment, 1972-82)</td>
<td>-.752</td>
<td>1.17*</td>
</tr>
<tr>
<td></td>
<td>(.493)</td>
<td>(.359)</td>
</tr>
<tr>
<td>Tenure *</td>
<td>.108*</td>
<td>-.106</td>
</tr>
<tr>
<td>ΔLog(Sector Emp.)</td>
<td>(.054)</td>
<td>(.070)</td>
</tr>
<tr>
<td>5-Year Technological Shock</td>
<td>.306*</td>
<td>-.027</td>
</tr>
<tr>
<td></td>
<td>(.125)</td>
<td>(.099)</td>
</tr>
<tr>
<td>Tenure *</td>
<td>.0002</td>
<td>.015</td>
</tr>
<tr>
<td>5-Year Technological Shock</td>
<td>(.011)</td>
<td>(.019)</td>
</tr>
</tbody>
</table>

Log-L                              | -536.4       | -243.7        |
Observations                       | 382          | 382           |

Note: The other covariates included in the model are the same as in Table 2. Complete results are available on request.

* - denotes significance at the .05 level, two-tailed test.
Table 5 -- Competing Risks Model for Recall and New Job Hazards, 1983 Data, with 1977-1982 Sectoral Employment Growth and 1982 Unemployment Rate Variables (variable coefficients; standard errors in parentheses)

Panel A: Using Sectoral Employment Growth from 1977 to 1982

<table>
<thead>
<tr>
<th>Variable</th>
<th>Recall Hazard</th>
<th>New Job Hazard</th>
</tr>
</thead>
<tbody>
<tr>
<td>Tenure at firm</td>
<td>.043*</td>
<td>-.025</td>
</tr>
<tr>
<td></td>
<td>(.013)</td>
<td>(.019)</td>
</tr>
<tr>
<td>(\Delta\log(\text{Industry/Geographic Sector Employment})) mean = .019</td>
<td>-1.15*</td>
<td>1.26*</td>
</tr>
<tr>
<td></td>
<td>(.592)</td>
<td>(.500)</td>
</tr>
<tr>
<td>Tenure * (\Delta\log(\text{Sector Emp.}))</td>
<td>.079</td>
<td>-.096</td>
</tr>
<tr>
<td></td>
<td>(.053)</td>
<td>(.065)</td>
</tr>
<tr>
<td>Log-L</td>
<td>-542.2</td>
<td>-246.0</td>
</tr>
<tr>
<td>Observations</td>
<td>382</td>
<td>382</td>
</tr>
</tbody>
</table>

Panel B: Using the 1982 Unemployment Rate

<table>
<thead>
<tr>
<th>Variable</th>
<th>Recall Hazard</th>
<th>New Job Hazard</th>
</tr>
</thead>
<tbody>
<tr>
<td>Tenure at firm</td>
<td>.089*</td>
<td>-.124*</td>
</tr>
<tr>
<td></td>
<td>(.025)</td>
<td>(.045)</td>
</tr>
<tr>
<td>Industry/Geographic Sector Unemployment Rate mean = .101</td>
<td>8.85*</td>
<td>-3.73</td>
</tr>
<tr>
<td></td>
<td>(2.90)</td>
<td>(2.59)</td>
</tr>
<tr>
<td>Tenure * Sector Unemployment</td>
<td>-.460*</td>
<td>.725*</td>
</tr>
<tr>
<td></td>
<td>(.213)</td>
<td>(.321)</td>
</tr>
<tr>
<td>Log-L</td>
<td>-539.4</td>
<td>-245.9</td>
</tr>
<tr>
<td>Observations</td>
<td>382</td>
<td>382</td>
</tr>
</tbody>
</table>

(continued)
Table 5 (continued)

Note: The other covariates included in the models are the same as in Table 2, except that the 5-year employment growth rate and 1982 unemployment rate, and their interactions with tenure, replace the 10-year employment growth rate and its interaction with tenure from Table 2. Complete results are available on request.

* - denotes significance at the .05 level, two-tailed test.