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Working Paper Number	2002-022D
Revision Date	October 2002
Citable Link	https://doi.org/10.20955/wp.2002.022
Suggested Citation	Friedberg, L., Owyang, M.T., 2002; Explaining the Evolution of Pension Structure and Job Tenure, Federal Reserve Bank of St. Louis Working Paper 2002-022. URL https://doi.org/10.20955/wp.2002.022

EXPLAINING THE EVOLUTION OF PENSION STRUCTURE AND JOB TENURE*

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November 4, 2005

*The authors wish to thank Stephanie Aaronson, Tim Cogley, Wouter den Haan, Douglas Fore, Rubén Hernández, Rob Dittmar, Alan Gustman, Richard Ippolito, David Laibson, Trish Pollard, Tara Sinclair, Robert Shimer, and many seminar participants for helpful comments. Partial funding was provided by TIAA-CREF, the University of Virginia's Bankard Fund for Political Economy, and the Federal Reserve Bank of St. Louis. Abbigail J. Chiodo, Kristie M. Engemann, and Ganesh Seshan provided excellent research assistance. The views in this paper are the authors' alone and do not reflect the views of the Federal Reserve Bank of St. Louis or the Federal Reserve System.

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Abstract

Current and expected job tenure have fallen significantly over the last two decades. Over the same period, traditional defined benefit pensions, designed to reward long tenure, have become steadily less common. This paper uses a contract-theoretic matching model with moral hazard to explain changes in pension structure and job tenure. In our model, a decline in the value of existing jobs relative to new jobs reduces expected match duration and thus the appeal of DB pensions. This explanation is consistent with observed trends and suggests an additional consequence of technological change that has not been closely studied. [JEL Classification: J32, J63, J65]

Keywords: Pension, Defined benefit, Defined contribution, Contracts, Job tenure

1 Introduction

In the midst of the economic boom of the 1990s, the *New York Times* suggested that "the notion of lifetime employment has come to seem as dated as soda jerks, or tail fins" (Kolbert and Clymer 1996).[1] Most data sets show that job tenure, especially of male workers, has fallen over the last two decades, and the Survey of Consumer Finances shows that expected remaining time on the job declined too. Consequently, total expected job duration of full-time employees in the SCF fell from 27.2 years in 1983 to 22.3 years in 2001 among men and from 22.5 years to 19.5 years among women. The decline for women suggests that their rising attachment to the labor force was tempered by a decline in attachment to a particular job.

Workers have also experienced a major shift in pension coverage since the early 1980s. Traditional defined benefit pensions, designed to reward long tenure, have become steadily less common, while defined contribution pensions, which are largely portable, have spread. The link between job tenure and pension trends has not been closely examined, but it offers insights about both phenomena. Analyzing this link allows us to bridge key gaps in the literatures on job stability

[1] We have appropriated this quote, with thanks, from Neumark, Polsky, and Hansen (1999).

and on the structure of compensation. First, we develop a matching model with endogenous job destruction that can explain the use of deferred compensation contracts and their connection to job duration. Earlier models of pensions typically did not incorporate uncertainty about job duration, nor make explicit the nature of the worker's outside option – both of which are formulated in matching models and crucially affect the value of tenure-based contracts. Earlier models of job matching rarely incorporated the use of deferred compensation. A group of recent papers has begun to analyze tenure-based contracts designed to deter on-the-job search;[2] this paper uses a model with a simpler form of moral hazard to highlight how changes in the economic environment alter the feasibility of such contracts.

Second, we discuss what kinds of shifts in the stochastic productivity process can explain observed trends in job tenure and pension structure. The model does not require a change in the mean productivity of new matches, which would be difficult to gauge. Instead, we focus on two less drastic possibilities: (i) an increase in the frequency of shocks that reduce the value of existing matches relative to new matches, or (ii) an increase in uncertainty about future productivity. Thus, the model provides possible explanations for the observed decline in job tenure, which few researchers have analyzed. It also offers a new, endogenous explanation for the decline in DB pensions that differs from the focus of other researchers on exogenous changes in pension regulation. The reversal in emphasis here suggests the possibility that regulatory changes *responded* to an underlying increase in worker mobility.

Third, we argue that new technologies have reduced the value of existing jobs relative to new jobs and possibly uncertainty in the manner which we hypothesize. We document recent patterns of technological change, job tenure, and pension structure that support the empirical implications of the model.

The paper is organized as follows. In Section 2, we discuss trends in job tenure and pensions and we argue that regulatory changes do not fully explain the shift in pension structure. In Section 3, we review past research on the functions of DB pensions, which may discourage moral hazard, motivate match-specific investment, and deter on-the-job search.

In Section 4, we develop a matching model and incorporate Lazear's notions of DB pensions. We show that a contract that defers compensation conditional on tenure, mimicking a DB pension,

[2] Burdett and Coles (2003), Stevens (2004), and Friedberg, Owyang, and Sinclair (2005).

elicits optimal effort. However, the contract may break down in the face of shocks to the output process which make it riskier to get bound into a long-term relationship.

In Section 5, we present empirical results. Because comprehensive data are lacking, we do not estimate the model developed earlier. Instead, we show that there is a strong empirical relationship between job tenure and pension structure; that the value of long-term jobs appears to have dropped; and that higher rates of technological change in industries are associated with lower job tenure and lower DB pension coverage.

In Section 6, we conclude by linking our results to other research on the nature of new technologies. Many of the phenomena identified in earlier studies on technological progress support our explanation behind a decline in the value of long-term jobs. Our results suggest a further consequence of technological change that has not been closely studied.

2 Background

In this section, we set the stage by presenting trends in job tenure and pension structure. We also discuss the structure of typical DB and DC pensions. Lastly, we contend that, while pension regulation has changed a great deal, it does not fully explain the observed trends in pension structure.

2.1 Trends in job tenure

We find that both current and expected remaining job tenure fell in the Survey of Consumer Finances (SCF).[3] Overall, total expected job duration fell significantly by 18% for men and by 13% for women.[4] Our theoretical and empirical analyses later on will draw links between the decline in both job tenure and DB pension coverage over the same period.[5]

[3] The SCF began in 1983 and surveyed a new cross-section every three years, offering the longest consistent information on pension coverage and expected remaining tenure; we omit data from 1986, which had an unusual sampling frame. The primary disadvantages of the SCF are that it is cross-sectional and that industry and occupation are reported at a very aggregated level (with only 6-7 classification codes). Other longitudinal data sets are not suitable; the HRS and NLS focus on particular age ranges that do not reflect the full shift in pension structure, while the PSID and SIPP have limited information on pension structure and job expectations.

[4] Early research did not confirm anecdotal reports, like the one cited earlier, of a decline in long-term jobs (Diebold, Neumark, and Polsky 1997). Since then, evidence has mounted of a decline in male job tenure in most data sets (Neumark, Polsky, and Hansen 1999; BLS 2000; Jaeger and Stevens 1999; Bernhardt, et. al., 1999) except the Survey of Income and Program Participation (Gottschalk and Moffitt 1999). None of these earlier papers used the SCF.

[5] Whether the decline in tenure is dominated by voluntary or involuntary mobility remains unclear. Our model does not distinguish between them, since all matches end endogenously. Moreover, it is irrelevant for our primary

Current job tenure. Average job tenure of male full-time employees aged 22-59 in the SCF fell from 9.2 in 1983 to 8.6 years in 2001. Average tenure of female full-time employees rose from 7.2 years in 1983 to 7.9 years in 1992 and then fell back to 7.1 in 2001. Tenure trends tended to flatten between 1995 and 2001 in the SCF and reversed for some subsamples in Table 1. However, residual tenure declined steadily from 1989 on if we control for business cycle effects.[6]

Male job tenure fell across the board by experience level, as shown in Table 1. Average tenure of men with 0-5 years of potential experience – those least likely to have DB pensions – declined significantly from 2.8 years in 1983 to 2.0 years in 2001. Average tenure of those with 6-15 and 16-25 years of potential experience declined significantly from 4.9 to 4.0 years and from 9.9 to 8.6 years, respectively. In results that are not shown, tenure fell for workers who attended college as well as those who did not. Changes in job tenure among women apparently reflect a combination of increases in labor force attachment early on and secular declines in job tenure later.[7] Tenure rose and then fell a little for those with 16 or more years of potential experience, while it tended to remain steady early on and then fell more (and statistically significantly) for those with less potential experience.

Expected remaining job tenure. The SCF also asked people how long they expected to continue working for their current employer – providing a direct measure of expected job duration, which is a key element of the model we present later. Expected tenure is noisier than actual tenure, especially for the least experienced workers, which is the smallest group. However, the series show generally significant declines as well – so the drop in current tenure in Table 1 reflects more than a one-time reshuffling of workers into new jobs. Among full-time employees aged 22-59, expected remaining tenure fell significantly for men from 18.0 in 1983 to 15.9 in 1992 and 13.7 in 2001 and for women from 15.3 in 1983 to 13.6 in 1992 and 12.3 in 2001.

Table 2 shows expected remaining job tenure by gender and years in the labor market. Declines are observed across the board by gender and experience over the entire period from 1983 to 2001,

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contention that tenure and pension trends are linked endogenously; an increase in either voluntary or involuntary job exits would reduce the appeal of long-term compensation arrangements.

[6] Similar patterns are seen in the CPS. We removed business cycle effects by regressing average tenure on the unemployment rate (either the contemporaneous rate or the average across two years) and analyzing trends in residual tenure.

[7] We will focus more on men than women in the rest of the paper, as is common in other research on job tenure because of confounding supply-side increases in women's labor supply.

and also from 1992 to 2001 when the questions were completely consistent across years.[8] Among men, expected tenure fell most for those with the least experience. For example, for those with 6-15 years of current tenure, it fell significantly from 21.8 years in 1983 to 18.7 in 1992 to 15.8 in 2001. It fell by less for men with current tenure of over 15 years. Again, changes in expected remaining job tenure among women reflect a combination of rising labor supply together with declining job tenure. For both men and women, expected tenure declined more for the more educated compared to the less educated.

Total expected job duration. Adding together current and expected remaining tenure yields an estimate of total expected job duration. For men, total expected tenure fell from 27.2 years in 1983 to 24.4 years in 1992 (a decline of 10.1%) and 22.3 years in 2001 (a further decline of 8.6%). For women, the total went from 22.5 years in 1983 and 21.5 years in 1992 (a decline of 4.4%) to 19.5 years in 2001 (a further decline of 9.5%).

2.2 Pension structure and trends

At the same time that job tenure has been declining, DB pensions have become steadily less common. We confirm this using data from the SCF: among full-time employees with a pension, 69% had a defined benefit (DB) plan and 45% had a defined contribution (DC) plan in 1983, while 39% had a DB plan and 80% had a DC plan in 2001. Overall pension coverage declined somewhat at the same time, from 67% of full-time employees in 1983 to 59% in 2001, as part of the general move away from deferred compensation. Later, we show that the DB pensions that remain seem to have declined in value as well. All of these trends, moreover, may be muted by persistence in compensation policies; workers are much more likely to get a new type of pension by changing jobs than by the employer altering pension features within a job.

2.2.1 The structure of pensions

Defined benefit pensions. DB pensions offer a defined payout to workers after they leave an employer. We can summarize its value in terms of pension wealth P_t, the actuarially discounted real present value of the worker's expected future pension benefits if the job ends in the current

[8] The decline in expected job tenure cannot be attributed to a decline in the expected retirement age. Both the expected retirement age and the proportion who said they would never stop working remained roughly flat over much of the period, and, if anything, rose after 1995.

year t. Pension wealth accrual is the discounted change in pension wealth $\frac{1}{1+r}P_{t+1} - P_t$, if the worker stays one additional year and then leaves.

The path of DB pension wealth accrual is typically characterized by sharp spikes. While the specific parameters of DB plans vary a great deal across employers, Figure 1 shows pension wealth accrual in a particular DB plan in 1992.[9] Spikes are generated at the vesting date and the early and/or normal retirement dates, depending on the particular pension formula. The late spikes in DB pension wealth thus discourage worker mobility for many years after a worker starts a job, while the negative pension accruals later on encourage retirement. Among older workers with a pension in 1992, median pension wealth was about \$200,000 if workers stay in their job until age 65. Allen, Clark, and McDermed (1988) estimated that the pension loss associated with switching jobs for the average worker aged 35-54 is approximately half a year's earnings.[10]

Defined contribution pensions. Accumulated employer and mandatory employee contributions to DC plans are a form of deferred compensation.[11] The accrual of this pension wealth is much simpler, though: contributions go into an account which earns a return, and the account is portable after vesting, which is often immediate. The resulting smooth path of DC pension wealth accrual shown in Figure 1 stands in stark contrast to DB accrual and makes it clear that DC pensions are generally tenure-neutral.[12]

It should be noted that, if long-term matches remain valuable, then employers could include tenure-based incentives in DC plans to a greater extent than they do. While around 85% of DB plans currently have cliff vesting at the maximum allowed period of five years, less than half of DC plans do. Similarly, employer contributions are based on tenure in only about 10% of DC profit sharing plans (Mitchell 1999). Recent developments have also reduced the use of tenure-based incentives in DB plans, with some traditional DB plans getting converted over the last 5-10 years

[9] Plan data were obtained from employers by the Health and Retirement Survey and have been slightly altered, as described in Friedberg and Webb (2005), to protect confidentiality.

[10] Gustman and Steinmeier (1993) pointed out, nevertheless, that pension wealth may be quite small at the start of a job. They argued that the primary effect of DB pensions is to deter mobility of longer-tenure, rather than new, workers. Extensions of our model to include investment in match-specific capital would sharpen the distinction between mobility incentives of shorter and longer-tenure workers.

[11] Employees are typically forbidden from withdrawing or borrowing against their plan balances (Mitchell 1999). Voluntary contributions by employees do not constitute deferred compensation but confer tax benefits that may not apply to other forms of saving. In any case, the tax treatment of DB and DC plans is similar; contributions are tax-deductible, returns accumulate tax-free, and income is taxable.

[12] Friedberg and Owyang (2002a) discussed other differences between DB and DC plans which do not affect mobility incentives.

into cash-balance plans, which are funded as DB plans but accrue pension wealth smoothly like DC plans (Coronado and Copeland 2003).

2.2.2 Regulation of pension plans

The government has frequently altered and tightened pension regulations since 1974. These changes have set funding standards for DB pensions, extended tax incentives for DC pensions, and constrained the structure of pensions in order to, for example, limit the extent to which they favor high-earning employees. Researchers have suggested several ways in which regulatory changes may have caused the shift away from DB pensions. None of these appear to fully explain observed trends in pension structure, however.

First, as pensions have become increasingly regulated, the costs of administering DB plans increased. However, the cost of administering DC plans rose at similar rates for all but the smallest plans (Ippolito 1995).[13] Second, Clark and McDermed (1990) claimed that some of the regulatory changes limited the extent to which DB plans can be designed as incentive contracts of the type we model later.[14] DB pension wealth can still accrue highly nonlinearly, though, as in the plan shown in Figure 1. Moreover, such a shift could have responded to, rather than caused, an increase in worker mobility, which might have raised concerns about workers losing out on expected future benefits. Third, Ippolito (2001, 2003) argued that regulatory changes in reversion taxes allowed companies to escape their DB pension obligations more easily than before, which undermined worker confidence and motivated the shift to DC pensions. Coronado and Copeland (2003) found that only about half of the S&P conversions which they examined were in a position to be influenced by reversion taxes, however.[15] Even where pension wealth is difficult to appropriate – in unionized and government jobs, for example (Ippolito 2003) – DB pensions are becoming less common and DC pensions more common in the SCF. Fourth, a countervailing effect arises from enhanced funding standards, which should increase the willingness of workers to accept DB pensions.

[13] Kruse (1995) concluded that rising administrative costs might explain some but not all of the decline in DB pensions between 1980 and 1986. Note also that enhanced tax incentives can explain why DC pensions have spread but not why DB pensions have disappeared, since a worker can (and many do) have both types of plans.

[14] Before ERISA established maximum vesting periods, for example, many DB pensions vested only at the normal retirement date.

[15] Moreover, a majority of the conversions increased pension liabilities to existing workers.

Thus, considerable evidence suggests that these regulatory changes fail to explain the entire shift in pension structure. Our focus on endogenous explanations provides a complementary perspective and even suggests that regulatory changes may have responded to an underlying increase in worker mobility.

2.2.3 Other evidence about pension trends

Several sources of evidence indicate that the evolution of pension structure has been associated with structural shifts in the economy. Employer-reported plan data show that workers have shifted from jobs that typically offer DB plans to jobs that typically offer DC plans.[16] Aaronson and Coronado (2005) found that both aggregate and industry-specific changes in pension structure in the CPS were important. In the SCF, we find that pension coverage shifted at varying rates by type of job and type of worker. Using analysis-of-variance, year main effects explain just under half (48%) of the over-time variation in DB pension coverage – so half of the decline occurred uniformly across types of jobs. Year-industry interactions explain 22%, indicating differing changes in DB pension coverage across industries; year-occupation interactions explain 13%, and year-education interactions explain 15%.[17]

Other details of pension trends suggest similar factors at work. Inequality in pension coverage across workers has grown, mirroring patterns in earnings inequality (Bloom and Freeman 1992, Even and Macpherson 2000) – a trend that is frequently attributed as well to structural changes in the economy. In support of the hypothesis we outline later, Coronado and Copeland (2003) and Aaronson and Coronado (2005) also found significant relationships between industry-level labor mobility and shifts in pension structure. There is mixed evidence about one additional factor. Aaronson and Coronado argued that the rise in labor supply of married women with children, who have shorter average tenure, also helps explain the decline in DB coverage. It is not clear, though, whether the sectoral choices of such women caused or responded to changes in the structure of

[16] Clark and McDermed (1990); Gustman and Steinmeier (1992); Ippolito (1995); Kruse (1995); Papke (1999). According to the second and third papers, for example, the movement of workers across jobs explains half of the shift in aggregate pension structure. We do not use employer-reported data in our later analysis because it lacks characteristics of covered workers.

[17] We used survey weights and controlled for age and age squared; employer size; education, occupation, industry, and union coverage; interactions of occupation with education and with industry; and gender and interactions with education, occupation, and industry. Friedberg and Owyang (2002b) described pension trends within industries and occupations.

compensation.[18] Also, we find that SCF industries and occupations (though highly aggregated) with greater gains in the share of women in employment did not experience greater declines in the use of DB pensions.

3 Theories of DB Pensions

Past theoretical research has sought to explain the incentive effects of DB pensions. The model we develop later builds on the idea that DB plans are designed to encourage optimal effort and longer tenure.

DB pensions as incentive contracts. In a series of papers summarized in Lazear (1986), Lazear developed models in which employers structure compensation to deter shirking by workers whose effort cannot be observed perfectly. A DB pension motivates effort by workers who do not want to get fired and lose their "bond".[19] We incorporate this motivation for pensions and then explicitly define the nature of uncertainty about job duration and of the worker's outside option – key elements determining the value of tenure-based contracts. Lazear (1983) argued that DB pensions also function as severance pay to encourage efficient retirement in models with rising wage profiles, another element of an incentive contract. We could extend our model to generate a rising wage profile if we imposed restrictions on the extent to which compensation could be deferred through the DB pension.[20] We chose not to include an explicit retirement motive, however, since our model generates an endogenous termination date.[21]

Other possible motives for DB pensions. An alternative theory is that DB pensions attract more productive workers, rather than elicit higher productivity (Viscusi 1985; Ippolito 1994). Empirical

[18] The authors implemented an interesting instrumental variables strategy, but the instruments are debatable. In industry-level regressions of changes in pension structure on, among other variables, changes in industry-level employment of women who are married, have children, etc., the instruments were lagged changes in the same variables. Pension coverage is highly persistent, though, so lagged changes in labor force characteristics may well have a gradual effect on, or gradually respond to, changes in pension structure.

[19] Related ideas appeared in Becker and Stigler (1974). Similarly, employers may offer DB pensions in order to recoup sunk costs of hiring, firing, or job-specific investments; or to discourage on-the-job search by workers seeking better offers (Burdett and Coles 2003; Friedberg, Owyang, and Sinclair 2005; Stevens 2004).

[20] Akerlof and Katz (1989) showed that, in the absence of up-front performance bonds, a rising wage profile alone, without a pension, is insufficient to deter shirking early in the career. On the other hand, wage tilt in Ippolito (1994) is necessary when the DB pension is too small to deter quits, though Ippolito (1991) found that wage tilt had no significant effect on job tenure, while DB pensions did.

[21] There is little evidence that a change in retirement motives generated the shift in pension structure, since pensions of older workers have changed much less than pensions of younger and shorter-tenure workers. If anything, the move away from DB plans may have increased firms' use of temporary early retirement inducements (Lumsdaine, Stock, and Wise 1990; Brown 2000).

tests to distinguish these motives have faced difficulties in resolving identification problems (Allen, Clark, and McDermed 1993; Even and Macpherson 1990). Moreover, an endogenous explanation for a declining use of pensions for screening requires a decline in the value of screening. This seems implausible given other labor market trends such as the growth in earnings inequality among workers with similar skills, which has been widely interpreted as an increased return to unobserved ability that should enhance the need to screen workers.

The observed link between unionization and DB pension coverage has led other researchers to focus on theories of union bargaining.[22] However, evidence cited earlier shows that the decline in unionized jobs does not explain a great deal of the shift in pension structure. Moreover, in the model we develop later, a decline in a worker's bargaining power has an ambiguous effect. It will make shirking more attractive and thus *increase* the value of the pension contract, though at the extreme it destroys the pension contract entirely because the contract can no longer deter shirking at all.

Motives for DC pensions. As the use of DB pensions has decreased, why have portable DC pensions become more popular? As we noted above, DC plans are much less likely to include vesting periods and other tenure-related features, so many of them simply constrain the path of consumption. A remaining possibility is that pensions have value as a vehicle for saving, perhaps because individuals have trouble saving adequately on their own or because the government wishes to increase savings rates. Savings-related motives may help explain the use of both DB and DC plans but not the tenure-related structure of DB plans nor the shift away from that structure.

4 A Model of Pensions

We develop an incomplete-contracting job-matching model that incorporates insights from past research on DB pensions. Matching models offer a rich representation of the labor market and of the effects of uncertainty which is absent from earlier models of pensions. While many search and matching models feature exogenous job destruction and focus on the rate and duration of unemployment, we emphasize endogenous job destruction, which motivates the use of pensions and determines the duration of employment.

[22] Freeman (1985) argued that unions give a stronger voice to older, less mobile workers who use pensions to appropriate rents from younger workers with higher quit rates.

In order to develop our arguments, we present a Nash bargaining model with moral hazard which builds on den Haan, Ramey, and Watson (2000, hereafter DRW).[23] As in DRW, moral hazard induces endogenous match destruction. We propose a pension-contract model that discourages moral hazard and eliminates inefficient match destruction, although we do not demonstrate that it is the only contract that would do so. After presenting the pension model, we discuss changes in the productivity process that would lead agents to abandon the pension contract.

4.1 The baseline model with moral hazard

Our model illustrates the inefficiency generated when unobservable effort on the part of the worker affects future match productivity. While we specify a simple form of moral hazard – low effort today destroys the continuation value of the match – we will indicate how it stands in for a richer model in which a worker decides whether to invest in match-specific capital that keeps the output distribution from drifting down, while skill-specific technological changes erode the stock of capital.

The matching market. A continuum of atomistic unemployed workers and firms who are searching in the labor market in a given period meet each other with probability λ.[24] The matched worker and firm i get an output draw $Y_{i,t}$ and decide whether to produce. If they do not produce, they return to the matching market next period. They will decide to produce if the output draw exceeds a threshold value R, reflecting the surplus from producing today and from the option to get another output draw and produce in future periods.

Production. Output Y is drawn from a distribution $F(y)$ which is the same for all new matches.[25] Thus, while agents are identical *ex ante*, matches are heterogeneous in their actual production draws. In each period, agents decide whether to continue producing or rejoin the labor market and draw their outside options. If the match breaks up, the worker and firm receive b^w and b^f from their contemporaneous outside option (with $b^w + b^f = b$) and expect ϕ^w and ϕ^f (with $\phi^w + \phi^f = \phi$) from re-entering the matching pool. If they produce, the agents split the match

[23] Valletta (1999) did not write down a model but discussed how extensions to a similar model by Ramey and Watson (1997) could help explain the decline in job tenure. Ramey and Watson modeled bilateral shirking in a pure contract-theoretic framework without search. Our approach extends the severance contract which they outlined to the search and matching model employed in DRW.

[24] Friedberg, Owyang, and Sinclair (2005) explore the role of tenure-based compensation when workers search while on the job.

[25] At this point, we assume a stationary distribution of Y and suppress the time subscript; later, we discuss the implications of nonstationarity.

surplus by Nash bargaining, with a share θ going to the worker and $1-\theta$ going to the firm.

In addition agents are subject to moral hazard; for simplicity, we limit consideration to moral hazard by the worker. A worker who shirks gains x^w this period but reduces future match productivity and thus the continuation value $g(R)$. We assume that shirking causes $g(R)$ to go to zero, so the match is severed.[26]

We can define the joint continuation value of the match as

$$g(R) = \beta \int_R^\infty (y + g(R))\, dF(y) + \beta \int_0^R (\phi + b)\, dF(y). \tag{1}$$

The continuation value equals the value of the match next period, discounted at rate β, if output exceeds the threshold value R, plus the discounted value of the outside option if output falls below R.

This allows us to define what the worker and firm gain from re-entering the matching pool as

$$\phi^w = \lambda\beta \int_R^\infty (\theta\,(y + g(R) - \phi - b))\, dF(y) + \beta\,(\phi^w + b^w) \tag{2}$$

$$\phi^f = \lambda\beta \int_R^\infty ((1-\theta)\,(y + g(R) - \phi - b))\, dF(y) + \beta(\phi^f + b^f). \tag{3}$$

These values depend on the probability of re-matching λ and subsequently drawing a satisfactory level of output (exceeding the threshold R) or alternatively remaining in the matching pool until the subsequent period.

Joint surplus from the match is defined as the value of the match less the value of re-entering the matching pool, $Y + g(R) - \phi$, and gets split according to the worker's bargaining share θ. This means that we can define the wage paid to the worker each period as the worker's portion of the surplus plus his outside option less his portion of the match continuation value. Under Nash bargaining, this is equivalent to the worker's share of output, so $w_t = \theta Y_t$.

Incentives in the presence of moral hazard. When the agents produce, the current value of the match is $Y + g(R)$, current output plus the continuation value of the match. This depends on the

[26] Suppose that match productivity is a function of match-specific human capital which must be kept current at a cost x^w to the worker. When the worker fails to update her specific human capital, match productivity falls enough to induce the firm to sever the match.

threshold output level R, which satisfies

$$R + g(R) = \phi + \max\{x^w, b\}, \tag{4}$$

where $\phi = \phi^w + \phi^f$ and $b = b^w + b^f$. When $Y_t = R$, agents are just indifferent between continuing or breaking up the match. If the moral hazard premium x^w exceeds the outside benefit b, then R rises by the difference, as we demonstrate below. The increase in R in the presence of moral hazard raises the expected value of the wage, which compensates the worker for forgoing the moral hazard payment.

We illustrate the impact of moral hazard on sustainable matches in Figure 2. $Y + g(R)$, match output plus the continuation value, appears on the vertical axis, and $\phi + b$, the outside option, appears on the horizontal axis. The Joint Productivity Threshold (Z) shows matches in which the firm's payoff (current period profit $Y - w$ plus continuation value g^f) exceeds its total outside option, while the worker's Incentive Compatibility (IC) constraint shows matches in which the worker's payoff (wage w plus continuation value g^w) exceeds the value of shirking plus the worker's outside option. IC lies a distance of $x^w - b$ above Z, since moral hazard imposes an additional requirement on current productivity to sustain the match. Matches below Z are jointly unproductive and are destroyed. Matches above IC are productive enough that the worker chooses high effort. Matches between IC and Z are broken up because workers choose low effort even though the matches are jointly productive. Essentially, $x^w > b$ creates a wedge between efficient and sustainable matches that require extra productivity in order to overcome shirking.

These scenarios are summarized in the following proposition.

Proposition 1 *Suppose that no steady-state displacements occur in the model without moral hazard (i.e., that $Y_t + g - \phi - b > 0$). For any $x^w > b$ and nondegenerate $F(y)$ with finite support in the model with moral hazard, the match is incentive compatible and thus is sustained if $w + g^w > x^w + \phi^w + b^w$, while the probability of match dissolution due to incentive incompatibility is strictly between zero and one.*

The incentive compatibility condition presented in the proposition requires that the worker's payoff (wage w plus continuation value g^w) exceed the value of shirking plus the worker's outside option. The proposition further implies that even though matches are jointly productive, there exists some Y for any x^w such that the match is not incentive compatible.[27]

An example. Suppose that the agents draw productivity Y from a standard uniform distribution. Then, in the solution to (4) that avoids the moral hazard problem, reservation output must satisfy:

$$R_{MH} = x^w + \frac{\beta(1-\lambda)\left(2b(1-R_{MH}) - (1-R_{MH}^2)\right)}{2k},$$

where $k = 1 - \beta(1-\lambda)(1-R)$.[28] If $x^w > b$, it drives up the minimum output R_{MH} required to sustain the match, changing the resulting values of g and ϕ. R_{MH} exceeds the reservation threshold computed in the absence of moral hazard,

$$R_N = \frac{2b - \beta\left(1-\lambda\right)\left(1-R_N^2\right)}{2k},$$

since higher productivity is required to deter the worker from shirking and sustain the match.[29]

The impact of moral hazard. Given the model (4), the worker will shirk if the value of not shirking and sustaining the match (the wage plus continuation value) is smaller than the payoff from shirking (the premium x^w and outside option). A higher value of θ, the worker's bargaining power and consequent share of future match rents, reduces the incentive to shirk. A higher λ, the probability of re-matching, raises the value of the outside option and hence the incentive to shirk. A higher b, the contemporaneous outside option, as long as $b < x^w$, increases the reservation threshold R_{MH} but by less than it would increase R_N. This is because match surplus, and therefore the wage and continuation value, continue to be determined by b, but R is now determined in part by x_w as well, so b has a reduced effect.

To understand the magnitude of the efficiency loss in response to some of these parameters, we

[27] A formal proof of a similar proposition appears in DRW. They show that for output below the reservation threshold, the derivative of R with respect to x is strictly positive. Thus, if no steady state dissolutions occur, x can be raised such that $R > 0$, so some matches that are dissolved in the MH model would not be dissolved in the N model.

[28] The solution is obtained by jointly solving the expressions for reservation productivity R, the outside option ϕ, and the continuation value g.

[29] A higher value of the output threshold R raises the continuation value g but also raises the value ϕ of re-entering the matching pool. R itself is a positive concave function of the contemporaneous outside option b. Additionally, for high values of b, R increases in the discount rate β and decreases in the probability of rematching λ. Lastly, $R_{MH} > R_N$ because $x^w > b$ and $0 < R < 1$.

compare aggregate output in period t in the moral hazard model, $\tilde{Y}_{MH} = \int_{R_{MH}}^{\infty} dF(y) = 1 - R_{MH}$, with aggregate output in the Nash model, $\tilde{Y}_N = \int_{R_N}^{\infty} dF(y) = 1 - R_N$. The productivity loss resulting from shirking is

$$\Lambda = \frac{\tilde{Y}_N - \tilde{Y}_{MH}}{\tilde{Y}_N} = \frac{R_{MH} - R_N}{1 - R_N}$$

which is always non-negative since $R_N < R_{MH} < 1$.

Figure 3 plots the productivity loss Λ, shown on the vertical axis, as the shirk premium x^w and the outside option b vary, given other reasonable parameter values ($\lambda = 0.3$, $\beta = 0.95$, $\theta = 0.5$). Since output per period lies between 0 and 1, we analyze values of x^w and b that are of the same order of magnitude. As we noted above, the productivity loss increases with x^w and b, since they make shirking more attractive. For example, the productivity loss ranges from 0 to 15% for $b = 0.66$ and x^w rising from 0.66 to 0.7, and it reaches as high as 50% when b and x^w exceed 0.9.

4.2 The pension model

Matches in the moral hazard model are vulnerable to incentives that raise payoffs to the worker today but destroy the future value of the match. This generates inefficient outcomes by forcing the dissolution of matches that are jointly productive. Here, we show that a deferred payment conditioned on match tenure – structured like a DB pension – can change the worker's incentives. The contract induces the worker to devote full effort and can be constructed to ensure that the match yields the same positive net productivity, so matches are efficient. We demonstrate that the pension contract remains incentive compatible and avoids inefficiency associated with moral hazard.

The pension contract. Suppose that the firm and worker write a contract $\{\overline{w}, W, T\}$ with the following elements:

- The worker collects wage $w = \overline{w}$ in each period when he is working and $t < T$.
- The worker collects $W(T)$, a lump sum, if he is still employed at time T.

Without loss of generality, we will set $\overline{w} = 0$ for the rest of this discussion.[30] We will discuss

[30] Enforcement considerations or risk aversion (as in Burdett and Coles 2003) would affect the actual tradeoff between $\overline{w}$ and W.

the choice of W and T later. We assume that the firm is prevented from breaking up the match if $Y_t > R_N$ and $t < T$. We also assume that the match breaks up if $Y_t < R_N$, even if $t < T$. Thus, the firm pays out W at time T as long as $Y_t > R_N$ each period. The assumption that firms are prohibited from severing productive matches but allowed to sever unproductive matches rests on the observability of Y_t and R_N. It is crucial, however; if a firm could not break up a match once a contract is in place, the worker would have no incentive not to shirk. Therefore, we must appeal to reputation effects or age discrimination laws which make it more difficult to fire older workers systematically than to lay off workers when output suffers. Empirical evidence over the period in which we are interested indicates that obvious breach of deferred compensation contracts by employers is infrequent.[31]

The worker's incentives. Under the pension contract, the worker's continuation value g_P^w depends on the wage contract $\{\overline{w}, W, T\}$. Again assuming $\overline{w} = 0$, then at the outset

$$g_P^w = \beta^T W(T). \tag{5}$$

We need to demonstrate several things about the worker's incentives in order to prove that the pension contract is feasible. First, if the worker accepts the contract in period 1, she will not sever the match later. The continuation value grows in later periods since the value is fixed but the worker discounts it less. Thus, by induction, she will not sever the match in any period $t > 1$ unless the productivity distribution shifts (which we have not allowed for yet) such that the worker's outside option grows relatively more valuable.

Next, we summarize in the following proposition the worker's incentive to shirk after accepting the contract, along with the worker's incentive to accept the contract at the outset:

Proposition 2 *Suppose the worker's payoff to shirking is x^w. Then, the worker will accept the contract $\{\overline{w}, W, T\}$ as long as the shirk premium satisfies $x^w < g_P^w + \overline{w} - \phi^w$. Specifically for the case $\overline{w} = 0$, if $x^w < g_P^w - \phi^w$ in each period, then the worker will choose high effort.*

[31] Pontiff, Shleifer, and Weisbach (1990) found relatively small gains from DB plan termination in firms experiencing hostile takeovers, and Gokhale, Groshen, and Neumark (1995) found mixed evidence that hostile takeovers in a small sample of eight firms led to reductions in extramarginal wages of workers. Cornwell, Dorsey, and Mehrzad (1991) found little evidence of opportunistic dismissal of pensioned workers in a nationally representative survey. Petersen (1992) found that DB plan termination was more likely in firms with more valuable implicit promises to workers, but that taxes had a greater effect on termination, while Coronado and Copeland (2003), as we mentioned earlier, found that a majority of cash balance conversions actually raised pension liabilities to existing workers.

Thus, the worker accepts the contract and does not shirk if the shirk premium is smaller than the value of the contract less the value of the outside option. We can check to see under what circumstances the contract in which the worker is paid $\overline{w} = 0$ until time T and W at time T satisfies the inequality. Substituting (5) for g^w and substituting for the value of ϕ_N^w we solved for above, this requires

$$x^w < \beta^{T-t} W(T) - \theta\beta \frac{b\left(k - \lambda(1 - R_N)\right) + \lambda J}{(1-\beta)\, k} \tag{6}$$

for all $t = 1, 2, ..., T$. As we mentioned above, the constraint is more likely to bind the lower is t. As time passes, the worker gets closer to the pension payoff and is less likely to shirk and risk getting fired.

Note that condition (6) determines the *minimum* W necessary to provide the worker with the proper incentives to ensure the match is both incentive compatible and yields joint net positive productivity. The actual choice of W could be modeled as depending on θ, the bargaining weight that determines the split of current-period surplus.[32]

Comparative statics. In order to understand how condition (6) governs feasible values of W and T, we analyze the impact of the threshold level of output R_N and then the fundamental parameters that determine R_N. (6) identifies the highest sustainable shirk premium x^w for a given conditional output $J = \int_R^\infty y dF(y) = \frac{1}{2}(1 - R_N^2)$ and severance risk $\pi = \Pr[Y_t < R_N] = \int_0^R dF(y) = R_N$. These two quantities are in tension as the reservation threshold R_N changes. Higher R_N reduces conditional output J and hence the value of re-entering the matching pool by reducing the likelihood that a productive match is formed; this raises the sustainable shirk premium for a given W. However, higher R_N also raises the severance risk π, so that staying matched becomes more uncertain; this reduces the sustainable shirk premium. At low R_N, the effect on J dominates the effect on π, making the required pension payoff W for a given termination date T relatively insensitive to changes in R_N. As R_N increases, π takes over and small changes in the reservation threshold have increasing effect on the sustainability of the pension contract.

Figure 4 shows how the minimum value of the pension W that satisfies (6) is affected by some of

[32] A possible choice of W is the present value of Nash bargaining each period, so $W = \sum_{i=1}^{T} \left(\frac{1-R_N}{\beta}\right)^i \left[\theta y_{t+i|t} - \bar{w}\right]$. This is also the expected future discounted value of the match less each period's wage payment, where R_N is the severance risk, $y_{t+i|t} = E\left[y_{t+i}|\Omega_t\right]$, and Ω_t is the information available when the contact is written.

the model's fundamental parameters. It shows how the minimum W, expressed as a percentage of the total expected value of the match at time T, changes as the shirk premium x^w and the vesting date T change, given other reasonable parameter values ($\lambda = 0.3$, $\beta = 0.95$, $\theta = 0.5$, $b = 0.5$). Again, since output per period lies between 0 and 1, we analyze values of x^w of the same order of magnitude.

It is apparent from Figure 4 that the promise of the future pension offering the worker a share of all future expected output has a powerful effect in deterring moral hazard. Thus, raising the shirk premium from 0.5 to almost 1 has little effect on the minimum required W for a given T. Figure 4 also shows the tradeoff between the term of the pension and its payoff; as noted above, an earlier termination date T allows for a lower payment W, given x^w. If x^w is 0.7, for example, a termination date of 25 periods requires a pension worth at least 27% of total expected Nash output, while a termination date of 15 periods requires a pension worth 10% of total output.

However, at sufficiently high values of x^w (approaching or exceeding 1, the maximum value of per-period output) the pension contract is no longer viable. The only way to deter shirking is to continue to increase W as x^w rises, but this is only profitable for the firm if it also extends T. Extending T raises the risk that the match will be severed before T, though. At some point, which is governed by (6), the firm cannot offer a high enough W to get the worker to accept the necessary increase in T, even if the worker's discount rate β gets very close to 1.

Summary. The contract $\{\overline{w}, W, T\}$ will be accepted by both agents and enhances efficiency when x^w satisfies (6). In the next subsection, we discuss how changes in the productivity process affect the pension contract.

4.3 Expected tenure and the productivity process

The previous subsection demonstrated how the DB pension (the lump-sum payoff W at time T) can resolve the inefficiencies resulting from moral hazard. In the model we laid out above, match productivity does not drift as it does in DRW, so the continuation value remains constant. The pension contract will also be effective if match productivity drifts upward, boosting the continuation value over time. In this section, we analyze the implications of other specifications of the productivity process – downward drift that reduces the productivity of existing matches relative to new matches, or an increase in uncertainty. Later on, we discuss the corresponding technology shocks which we have in mind.

Changes in the stochastic productivity process of the type we discuss here will reduce expected job tenure, so it is useful to define worker's expected tenure as $E(\tau) = \frac{1}{1-R}$. The recurrent theme is that, if the decline in job tenure is severe enough, it will render the pension contract infeasible, since the worker no longer accepts deferral of payment because the risk of exogenous separation becomes too high.

Downward drift in the productivity of existing matches. Initially, we considered only matches which were jointly productive. Suppose now that output each period is drawn from successively less favorable probability distributions, so $F_{t+1}(y) > F_t(y)$ for all t. This implies a time-dependent continuation value in which $g_{t+1}(R) < g_t(R)$. The Nash bargaining model then implies an increasing reservation productivity $R_{t+1} > R_t$, since conditional output $J > 0$ for all y; only a higher draw will make the agents willing to continue the match in the face of worsened long-term prospects.[33]

The resulting condition $R_{t+1} > R_t$ has implications for job tenure. The severance risk $\pi = \int_0^{R_t} dF_t(y)$ increases when either the distribution becomes less favorable or reservation output rises, increasing the likelihood of separation. This lowers expected job tenure and thus the expected value of the pension to the worker, since the probability that the match lasts until T declines. As we noted at the end of the previous subsection, this effect will reduce the maximum sustainable shirk premium, and at some point the contract breaks down. Put differently, as the likelihood of exogenous separation rises, the payoff date in the contract must get increasingly close to the initiation date for the worker to accept the risk of exogenous separation. However, reducing T also reduces the nominal value W of the pension which the employer is willing to offer. At some point expected tenure $E(\tau)$ becomes small enough that the worker will not accept the contract. Consequently, a decline in expected tenure will lead to fewer and less valuable pension contracts.

Increased uncertainty in the productivity process. The productivity threshold R is unaffected by a change in the variance of the productivity process. Hence, a mean-preserving spread in the productivity distribution raises the probability that the match will fall below the cutoff value R at some future date, if R lies below the mean of the productivity distribution.[34] Again, the terminal

[33] In the type of human capital model we have alluded to, these shocks can result from the introduction of a new technology which erodes the value of existing skill-specific human capital.

[34] The implications of a mean-preserving spread are reversed if R lies above the mean, but this case seems unlikely as it requires that the mean productivity draw is insufficiently high to warrant preserving matches.

date T must be reduced for workers to accept the pension, but that reduces the nominal value W of the pension which undermines the value of the pension to workers.

Summary. The preceding discussion provides intuition as to the breakdown of DB pensions. Matches with stable or increasing continuation values can benefit from deferring payment to the worker in order to provide incentives that are unavailable in a standard Nash bargaining model. These contracts preserve jointly efficient matches that would ordinarily be severed. Matches with decreasing continuation values, however, might not sustain the pension contract. Additionally, a mean-preserving spread in the distribution of future productivity draws may sufficiently raise uncertainty about match duration such that the pension contract cannot be sustained.[35]

4.4 Government regulation in the pension model

In the moral hazard model, efficiency is enhanced by moving from the standard Nash contract to the DB pension contract if expected match tenure is sufficiently high. We have focused on changes in the productivity process that undermine such contracts, but it is also possible that they are undone by government regulation. Suppose the government dislikes the outcome that some workers suffer exogenous separation before they collect their pension and requires that workers be guaranteed their accrued pension wealth if matches end before T. This destroys the firm's ability to influence worker effort.

We can evaluate the loss resulting from rekindling the moral hazard problem based on our earlier definition of the efficiency loss $\Lambda = \frac{R_{MH}-R_N}{1-R_N}$ arising from moral hazard in the absence of pensions. Figure 3 showed how the efficiency loss Λ increases as the shirk premium x^w and the outside option b rise, given other reasonable parameter values. For values of $x^w$and b around 0.5 (recall that output draws are bounded between zero and one), the efficiency loss can reach 8%, while for values around 0.65 to 0.7, it can be twice as high. Thus, our model presents the policymaker with a choice between social efficiency versus helping a fraction of workers who experience bad luck by mandating portability of pensions.

[35] One must consider other possible contracts at this point. Ramey and Watson (1997) showed that contracts with severance payments or punishments can sustain matches in the efficient but incentive-incompatible region. However, such contracts are rare, perhaps because they are not easily enforced or yield socially inefficient litigation upon separation.

5 Empirical Evidence

In this section, we present empirical evidence that supports the hypotheses we have presented. It is difficult to estimate our model directly, given the absence of linked employee-employer longitudinal data or even employee longitudinal data with details about the structure of compensation and productivity. The alternative is to test implications of the model that relate to pension structure and job tenure. There are several types of evidence from the SCF and the CPS that we bring to bear.[36]

First, we show that job tenure is related to pension structure in the way that our model presumes. We find that workers with a DB pension and with more valuable DB pensions have longer tenure than workers with DC pensions or workers with no pensions. Second, we show outside evidence that the value of long-term jobs has dropped, since DB pensions have declined somewhat in value and estimated earnings-tenure profiles have flattened out considerably. Third, we present evidence that links technological progress to both the structure of compensation and job tenure. Industries with higher rates of computer use and overall investment have lower rates of DB pension coverage and lower average job tenure, and further these relationships were more negative in the 1990s than in the 1980s. If, instead, government regulation induced the shift in pension structure, there would be no reason to expect such links.

5.1 Pension structure and job tenure

We show that workers with DB pensions have longer current and expected total job tenure than both workers without pensions (as in Allen, Clark, and McDermed 1993) and workers with DC pensions (in contrast to Gustman and Steinmeier 1993). While we view pension type as the best proxy for the degree to which compensation is tied to tenure, we find further that workers with more valuable DB pensions have longer tenure, controlling for the level of earnings. Note that we do not estimate a structural model of compensation and mobility, so our approach does not

[36] We noted earlier the reasons we concentrate on tenure and pension information from repeated SCFs and on tenure information from repeated CPSs. We also use some data from the April 1993 CPS, the last time questions were asked about pensions. The notes to each table describe the sample, definition of variables, and estimation details. When using SCF data in this section, all coefficient estimates and standard errors are computed from regressions run on multiple implicates (Rubin 1987).

distinguish whether DB pensions cause longer tenure.[37]

Regressing job tenure on pension type. We ran several regressions, separately for men and women, using both the SCF and the last pension supplement of the CPS.[38] In the SCF results shown in Table 3, all of the pension variables have statistically significant effects on job tenure. Male workers with a DB pension have been in their jobs about 5 years longer than workers without a pension, depending on the specification. Female workers with a DB pension have been in their jobs about 4 years longer. Workers with both types of pensions have been in their jobs about half a year longer than workers with only a DB pension, but the difference is generally not statistically significant. In comparison, workers with a DC pension have been in their jobs 2-3 years longer than workers without a pension, significantly shorter than workers with a DB pension. While job characteristics such as industry and occupation may explain both pension structure and job tenure, including such controls in the second and fourth columns reduces the estimated effect of pensions on tenure by a year or less.

The relationship between pensions and job tenure remains strong when year effects are included in the third and fourth columns, so it does not reflect a spurious correlation between two trending variables. Moreover, it persists if we interact pension type with year, in results that are not shown.[39] Therefore, we can conclude that the same workers who are experiencing a decline in DB pension coverage are spending less time in their jobs. Although we do not ascribe a structural interpretation to the magnitude of the estimated effect, the observed decline in DB pension coverage between 1983 and 1998 is associated with a decline in current job tenure of 0.9 years for males and 0.6 years for females, according to regressions (4) and (8); this is of the same order of magnitude as the observed decline in tenure.

[37]As mentioned earlier, we were unconvinced by previous attempts to estimate endogenous selection into jobs with DB pensions. While including earnings on the right-hand side in our regressions is just one possible source of endogeneity, it furthers our goal of describing the empirical relationship between tenure and the *structure* of compensation, while controlling for the level. A caveat to these results is that workers report pension type with considerable error (Gustman and Steinmeier 1999); yet, Chan and Stevens (2004) found that retirement behavior responds to misperceptions about pension incentives.

[38]We run all tenure regressions separately by gender because men and women exhibited different secular trends. It would be preferable to run regressions on the job exit hazard rather than on current tenure, which represents incomplete spells, but the mean of complete and incomplete job spells will be the same if spell length is not duration dependent. If it is, then a linear regression on tenure can be viewed as a first-order Taylor expansion of more complicated specifications (Freeman 1980). We obtained similar results from regressions on the log of tenure, as in Even and Macpherson (1996).

[39]The estimated response to DB pensions declined by 2 years or less between 1983 and 2001 and remained significantly higher than the response to DC pensions throughout the sample period.

We ran additional regressions with similar results that are not shown here. We used total expected job duration (the sum of current and expected future tenure) as a left-hand side variable. Workers with a DB pension have total expected tenure that is 5.0-7.0 years longer than workers without a pension, while workers with a DC pension have total expected tenure that is 2.5-4.0 years longer. We estimated almost the same relationship between pension structure and job tenure in the April 1993 CPS.[40]

Regressing job tenure on the value of DB pensions. We used data reported in the SCF to compute the value of DB pensions and added that information to regressions like those reported in Table 3. In some regressions, we include information which individuals with DB pensions report about the pension benefit they expect to receive if they stay in their jobs as long as intended. Because that is endogenous with expected tenure and reported with error, in other regressions we include the average benefit imputed on the basis of earnings, industry, occupation, education, unionization, employer size, and gender. This is similar to the approach in Gustman and Steinmeier (1993), described in more detail below, of including an imputed measure of pension backloading.

We find some interesting results in Table 4. First, a higher value of one's DB pension at retirement is associated with significantly longer tenure. The semi-elasticity of tenure with respect to the log monthly pension benefit (the statistic reported by most of the post-1983 sample) is around 0.6 when self-reported information is included (in regressions labeled a) and 0.3-0.45 when the average benefit is included (in regressions labeled b).[41] This implies, for a male with the median value of expected future pension benefits ($958 per month in 2001 dollars) and according to (4a), that job tenure is half a year longer than for someone with the 25th percentile value ($400) and almost a year shorter than someone with the 75th percentile value ($2076).

Second, once we control for DB pension value, then the additional effect of having a DB pension is reduced, compared to Table 3. It lies in a range between 0.5-3.5 years (versus 4-6 years in Table 3), in some cases not statistically distinguishable from zero and in others not statistically distinguishable from the effect of DC pensions, which remains in the range of 2-3 years. Together, these two findings support our hypothesis that DB pensions are used to extend job tenure, since

[40] These coefficient estimates are reported in Friedberg and Owyang (2002b). It is noteworthy that the estimated effect of pensions declines by only a couple tenths of a year when detailed industry and occupation codes reported in the CPS are included.

[41] The numbering of the regressions in Table 4 parallels that in Table 3. When we use total expected tenure on the left-hand side, the estimated coefficients are about twice as large.

the differential effect of DB pensions on tenure is operating through the value of the pension.

Comparison to other results. The results in the regressions described above, duplicated in two data sets, differ importantly from Gustman and Steinmeier (1993). They found similar mobility rates for workers with DB and DC pensions, perhaps in part because of differences in the time period and in the measurement of key variables in their data. They used the SIPP, which is a panel and allowed them to focus on mobility rather than tenure.[42] However, they used data from only 1984-85, before DC plans became common. Also, in their words, "the SIPP question sequence on plan type is atypical" (p.303, 1993) and overstated the prevalence of DC plans.

Another difference was that they added a selection-adjusted imputed measure of the compensation available in alternative jobs on the right-hand side of their mobility equation. This is more ambitious than our estimation approach. Their results suggested that pensioned workers faced *worse* alternatives relative to their current jobs than did non-pensioned workers and that the compensation differential had a greater effect than pension wealth in deterring mobility. However, this result hinges on knowing the terms of jobs available to workers who do not move, though no motivation was offered for the identifying exclusion restrictions used in predicting mobility and imputing alternative compensation.[43] Lastly, they found another anomalous result that we did not – in their estimates, pension backloading had a much greater effect in deterring mobility than current compensation, but in our estimates the semi-elasticity of tenure with respect to earnings is somewhat larger (in the range of 1.5-2.5 years) than the semi-elasticity with respect to pension value.

To sum up, their evidence that the alternative compensation premium, rather than the structure of DB pensions, explains the pension-mobility relationship is not convincing, in our view. In contrast to their results for a particular data set and year, we find evidence of a robust long-term relationship between pension structure and mobility.

[42] The reliability of job mobility data in the SIPP is unclear, since it does not show the decline in job tenure which is apparent in other surveys, as we discussed earlier.

[43] Age, marital status, children under 18, and home ownership were included in the mobility equation but excluded from the current and alternative compensation equations. However, these variables are correlated with wages in simple OLS regressions using our data, probably because they reveal something about unobservable productivity. Pension coverage was included in the compensation equations but excluded from the mobility equation.

5.2 The value of long-term jobs to workers

While we do not have data to estimate changes in the value of a long-term job, we have some evidence about the value to workers. We show two ways in which tenure-related compensation appears to have shrunk: DB pensions appear to have lost value, and the degree to which earnings rise with tenure has fallen a great deal. After presenting these results, we will discuss how they can be interpreted in light of our model.

The value of DB pensions. As described above, from 1989 on the SCF reports the benefit people expect to receive when they leave the firm.[44] In order to detect changes over time, we regress this variable on year dummies. However, the expected pension benefit depends not only on the degree to which a DB pension defers compensation but also on the worker's expected tenure, so we include detailed controls for current and expected remaining job tenure. We also control for current earnings, in case overall compensation declined, and in some cases we control for other individual and job characteristics to isolate shifts in the terms of particular jobs, rather than shifts in the composition of jobs.

We find mixed results, with somewhat high standard errors. With the caveat that they may to some extent reflect changes in tenure, the regressions in Table 5 show that DB pensions declined substantially in value between 1989 and 1998, while increasing somewhat in 2001, though not significantly. For the average male with a DB pension, according to the results in specification (2), the expected monthly benefit declined steadily from 1989 through 1998 for a statistically significant total of $352, which compares to the average in 1989 of $1801. The expected monthly benefit then rose by $150 between 1998 and 2001, and the total decline of $202 was not statistically significant. Among females as well, according to specification (4), the changes tended to be negative but non-monotonic. The overall decline of $294 by 1998 was, again, statistically significant and compares to an average 1989 benefit of $1091, while the decline of $246 by 2001 was not statistically significant.[45]

The relationship between tenure and earnings. A common practice in the labor literature is to estimate a "return to tenure", with current earnings on the left-hand side and tenure and

[44] While the SCF asked many questions about pensions, we are hesitant to use self-reported information at any greater level of detail, given the lack of familiarity most people display about specific features of their pensions.

[45] Based on the estimates in Table 5, we verified that the decline in the value of DB pensions, if it occurred for exogenous reasons, is not nearly large enough to explain the overall decline in job tenure.

other measures of human capital on the right.[46] If we find that the tenure premium has fallen, it suggests that long-term jobs have become less valuable. Farber (1999) described the problem of interpreting such estimates, since many theories (including ours) predict that compensation is structured to influence tenure. He argued that such estimates are interesting, nonetheless, in revealing the nature of firm-level compensation structures. This motivates our analysis as well, and later we interpret the results in light of ours and other possible models. We used data from CPSs between 1983 and 2002, since the CPS offers large sample sizes and, importantly, detailed industry and occupation controls.[47] We estimated log earnings equations for men and women separately and included quartics in tenure for each CPS year, while controlling for the level of earnings and allowing the effects of several other covariates to differ across years.

Table 6 shows the earnings premium paid to average male and female workers with 5, 10, 15, 20, and 25 years of tenure, compared to a worker beginning a job. The results support our claim that the value of long-term jobs declined. After some gains in the 1980s, the tenure premium dropped sharply after 1987 for men and after 1987-91 for women. Notably, it continued to drop between 1996 and 2002 while actual tenure tended to flatten out, as we noted earlier. For example, the earnings premium enjoyed by males with 10 years of tenure rose from 21.3% in 1983 to 27.6% in 1987 and then dropped to 14.1% in 2002.[48] The overall decline in the tenure premium between 1983 and 2002 was statistically significant, and it fell the most for males with 15 years or fewer of tenure. The drop-off for females occurred a little later but was sharper – the earnings premium rose from 24.0% in 1983 to 29.4% in 1987 and then fell to 11.4% in 2002 for females with 10 years of tenure.

Balan (2003) found an earlier decline in the tenure premium for both men and women in the PSID. Using the panel data to instrument for job tenure, he estimated a significant drop of roughly 3/4 of a percentage point per year between 1981 and 1992 among private-sector non-unionized male

[46] While our model does not feature a rising wage profile, we alluded earlier to extensions in which a rising wage is part of an incentive contract to lengthen job duration.

[47] We do not use earlier CPS tenure supplements because the wording of key questions changed. We adjusted the reported tenure data for half-year rounding among those with 1-2 years of tenure, though in a simpler way than Diebold, Neumark, and Polsky (1997). We did not adjust the data for heaping at five-year intervals, as they did; in their 1996 paper, they showed that adjustments for rounding and heaping did not affect conclusions about the magnitude of job tenure trends. When we tried adjusting the sampling weights, as they did, for differences in nonresponse to the tenure question by age, sex, and race, the results were virtually identical.

[48] As a point of reference, Topel (1991) estimated that the selection-corrected earnings premium for males with 10 years of tenure was over 25%.

workers, with a slightly greater decline for recent entrants to the job market.

In additional results, we find that the tenure premium dropped in industries with declining tenure, supporting the idea that these trends are linked. We regressed the median real tenure premium on average job tenure in the same industry and found that a one-year decline in average job tenure is associated with a significant 2.3 percentage point decline in the tenure premium.[49]

Interpretation. Our results show that both DB pensions and the earnings premium associated with longer tenure have shrunk in value. What can we infer from this about the value of long-term jobs? Since we control for the level of current earnings, these changes are not a consequence of an overall reduction or redistribution of match surplus but are related specifically to the duration of jobs.

Our model illustrated the tradeoff between the term and the value of the pension contract – as the likelihood of exogenous separation increases, the worker demands an earlier payoff, but the size of the pension that the firm is willing to offer falls and at the limit the viability of pensions is threatened. Thus, a decline in the value of remaining DB pensions supports our hypothesis that the value of long-term jobs fell. If increased regulation explained the shift away from DB pensions, it is not clear why it would also reduce the value of remaining DB pensions, while if anything tenure profiles might be made steeper.

The literature offers a number of explanations for a rising tenure premium, and the inferences we might draw from observing it decline are generally consistent with our main arguments. As noted earlier, a rising tenure profile may itself be a component of a tenure-based incentive contract. In this case, a reduction in the value of tenure-based contracts would not only flatten the tenure profile, but it would also reduce selective mobility, a more general explanation for observing a tenure premium whereby matches end selectively when their productivity draws fall below a reservation level. A final explanation is that workers are paid their marginal product and match-specific productivity rises with tenure, so a decline in the tenure premium would result from declining productivity of long-tenured workers – also a sign that the value of long-term jobs may have declined.

[49] We computed tenure premia and average job tenure in 45 two-digit industries that employed at least 100 people in each CPS survey. The regressions are weighted by the number of people from the CPS in each cell. If we control for year and industry effects (accounting for economy-wide changes and industry-specific values of the tenure premium), then the correlation is a little greater.

5.3 Technological change, pension structure, and job tenure

We have shown that both actual and expected job tenure fell and that deferred compensation has shrunk, from which we infer that the value of long-term jobs has declined. In our model, this will occur if there is an acceleration of shocks that erode the productivity of existing matches relative to new matches, or simply an increase in uncertainty about the future productivity of matches. We hypothesize that a shift in the nature of new technologies has had such effects.

As above, and in research by others on the nature and impact of new technologies, there are no data to test this directly. However, we demonstrate that some commonly used measures of technological change are related to both pension structure and job tenure as we hypothesize. We use data on computer use in jobs, which offers a straightforward way to measure the diffusion of a major new technology.[50] We also use data on investment, the capital stock, and total factor productivity (TFP) growth, standard measures of both embodied and neutral technological progress used in the macroeconomics literature.[51] We use industry-level data on these measures and match them to industry averages of job tenure from repeated CPSs and pension structure from the April 1993 CPS.[52] As above, we include average earnings in the regressions in order to detect changes in the structure of compensation while controlling for the level.[53] The April 1993 CPS also included information on firm size, which can control for possible scale economies in offering DB pensions as well as adopting new technologies.

Technological change and pension structure in industries. The only data source reporting both pension structure and earnings in detailed industries is the April 1993 CPS. We regress average DB pension coverage by industry on each measure of technological change and report the results in Table 7. We find that computer use has a negative relationship with DB pension coverage (in the upper left panel), and it becomes significant when we control for firm size – so larger firms

[50] In October 1984, 1989, 1993, 1997, and September 2001, the CPS asked individuals whether they used a computer at work. We compute average computer use in 50 industries in each year. In 2001, we also include as computer users those who reported that they had a home computer and used it for work.

[51] We use investment, capital, and TFP data from the Jorgenson Total Factor Productivity Series, which covers 21 manufacturing sectors and 14 other non-manufacturing sectors annually from 1959 to 1996; the manufacturing sectors are quite disaggregated, but other sectors are not. While we average the data over 10-year periods, we obtained generally similar estimates using averages over 5 years.

[52] These were measures used by Autor, Katz, and Krueger (1998) and Bartel and Sicherman (1999) to determine the impact of technological change on workers. It is much less useful to try to match these measures of technological change to SCF data, since SCF industry codes are so highly aggregated.

[53] Also, we weight regressions by the number of workers in each industry so the results reflect economywide averages.

had higher rates of both computer use and DB coverage. Earlier changes in computer use do not have a significant coefficient. TFP growth, investment, and capital all have significant negative relationships with DB pension coverage (in the lower left panel), and the estimated coefficient is hardly affected when controlling for firm size.

Based on the most detailed specification in column (3), an industry with a one standard deviation higher rate of computer use has a 4.0 percentage point lower rate of DB pension coverage, which is 8.3% of the mean rate. Based on the specification in (7), an industry with a one standard deviation higher rate of TFP growth (investment) has a 3.9 (4.2) percentage point lower rate of DB pension coverage. Note that the impact of observed variation in all of these measures of technological change is quite similar.

Technological change and job tenure in industries. Table 7 also reports the relationship between job tenure in the April 1993 CPS and measures of technological change. Computer use, investment, and the capital stock all have a significant negative relationship with job tenure. Adding firm size dummies again increases the estimated coefficient on computer use (in the upper right panel), though less dramatically than it did in the DB pension regressions. It raises the estimated coefficient on investment as well (in the lower right panel), though it increases the standard error considerably, while it substantially reduces the unexpected positive (and highly insignificant) coefficient on TFP growth. Based on the specifications in (6) and (1), a one standard deviation higher rate of computer use is associated with 0.87 years less in average job tenure, or 10.0% of average tenure, and a one standard deviation higher level of investment is associated with about 0.6 years less.

We also use repeated CPS job tenure supplements to explore the relationship over time with these measures of technological change, in Table 8. Computer use has a negative, significant relationship with job tenure in specifications (1) and (3); however, it shrinks a little and loses significance when controlling for year effects in job tenure in (2) and (4), and the estimate is smaller than it was in the April 1993 regressions. Investment and the capital stock also have a negative and significant association with job tenure in (7) and (9); the coefficients only decline slightly when year effects are included in (8) and (10), and they remain a little larger than the estimate in the April 1993 regressions.

Interestingly, we find evidence that these relationships changed over time when the sample is split before and after 1990. All the relevant coefficients are smaller and/or more negative in the later sample. Comparing (5) and (6), computer use has a negative but insignificant relationship with job tenure before 1990 and a more negative and now significant relationship after 1990, even when year effects are included. Similarly comparing (11) and (12), the relationship between investment and job tenure becomes more negative and significant after 1990, while the unexpected positive effect of TFP growth on job tenure disappears.

Summary. We find a consistent negative association between measures of technological change and industry-level averages of DB pension coverage and job tenure. Aaronson and Coronado (2005) found a similar relationship, using other data and other measures of technological change, strengthening our claim that the nature of technological advances has changed in a way that undermines the value of long-term jobs.

6 Conclusion

In this paper we have specified a model of DB pensions and job tenure. DB pensions eliminate inefficient job destruction resulting from moral hazard; in more complex models, the moral hazard can take the form of searching on-the-job or failing to invest in job-specific capital. The use of DB pensions is undermined, however, if expected job tenure declines. We have shown in this paper that both actual and expected job tenure fell along with the use of DB pensions.

We also used the model to demonstrate the types of changes in the stochastic productivity process which reduce expected job tenure and hence the use of DB pensions. In particular, we focused on shocks that increase uncertainty about future match productivity. To that end, we showed that industries which have experienced more rapid growth in technological progress also experienced greater declines in job tenure.

These results complement a large body of research on the shifting nature and pace of technological changes, suggesting that they have had the effects on jobs that we have in mind. Researchers have found that the diffusion of new, especially information-related, technologies has had a powerful effect on the level of compensation by raising earnings inequality across jobs (Gottschalk 1997, Acemoglu 2002). Our hypothesis is that it has altered the structure as well as the level of compensation.

The key reason for rising inequality, according to this literature, is that new technologies are largely skill and ability-biased. The average skill level of workers rose more in industries that experienced higher rates of investment in general and of computerization in particular (Autor, Katz, and Krueger 1998). Case study evidence suggests further that new technologies require not just greater but also new skills. Computer use is, obviously, one of the new skills; employers and individuals continue to devote substantial resources to computer training, even while computers have grown easier to use over time.[54] Besides that, computers have automated routine tasks while altering and often making more complex the performance of non-routine tasks (Levy and Murnane 1996, Autor, Levy, and Murnane 2002). Computerization has brought on further changes in required skills, workplace organization, and the delivery of services, requiring substantial training and other adjustment costs (Bresnahan, Brynjolfsson, and Hitt 2002).

Our evidence adds to this literature by suggesting that jobs have been reorganized in ways that loosen the ties of long-term relationships between workers and firms. Moreover, since both expected remaining job tenure and the use of DB pensions in new jobs has declined, our results indicate that the changes in the productivity process are permanent.

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Table 1: Current job tenure

	Men				Women			
	Average, by years of potential experience							
	0-5	6-15	16-25	26-35	0-5	6-15	16-25	26-35
1983	2.8	4.9	9.9	14.2	2.3	4.9	7.7	9.9
1989	2.2	4.8	8.9	14.7	1.9	4.0	7.6	10.5
1992	2.3	4.8	8.1	14.1	2.1	4.8	8.2	11.8
1995	2.0	4.6	8.5	12.9	2.1	4.4	8.4	10.7
1998	1.7	4.4	8.6	13.6	1.7	4.0	7.4	11.1
2001	2.0	4.0	8.6	13.6	1.9	3.9	7.1	10.7
change 2001-83	-0.8	-0.9	-1.3	-0.6	-0.4	-0.9	-0.6	0.8
(standard error)	(0.3)	(0.2)	(0.4)	(0.7)	(0.2)	(0.3)	(0.5)	(0.8)

Data source: Survey of Consumer Finances from 1983, 89, 92, 95, 98, 01. Respondents were asked "How many years in total have you worked for this employer?"
Sample: Full-time employees aged 22-59, except those who reported tenure that exceeded potential experience plus two years (about 1.5% of the sample).
Details: Means and standard errors are computed from multiple implicates (Rubin 1987), using survey weights. Years of potential experience is defined as age minus years of completed education minus six.

Table 2: Expected remaining job tenure

	Men				Women			
	Average, by years of potential experience							
	0-5	6-15	16-25	26-35	0-5	6-15	16-25	26-35
1983	18.3	21.8	20.2	13.2	15.3	18.3	17.9	11.4
1989	11.0	18.4	17.8	11.6	8.7	13.2	14.6	9.8
1992	16.7	18.7	17.5	11.9	13.6	15.8	14.9	10.9
1995	11.1	17.3	16.3	11.4	8.7	15.4	13.6	10.2
1998	10.6	16.1	16.3	11.0	10.1	13.0	14.7	10.4
2001	12.0	15.8	15.4	11.2	12.9	14.6	13.4	10.2
change 2001-83	-6.3	-6.0	-4.7	-2.0	-2.3	-3.7	-4.5	-1.3
(standard error)	(2.7)	(1.1)	(0.8)	(0.5)	(2.6)	(1.2)	(0.9)	(0.6)
change 2001-92	-4.7	-2.9	-2.1	-0.7	-0.6	-1.3	-1.5	-0.7
(standard error)	(3.4)	(1.2)	(1.2)	(0.6)	(5.0)	(1.8)	(1.1)	(0.6)

Data source: Survey of Consumer Finances from 1983, 89, 92, 95, 98, 01. Respondents were asked "How many years do you expect to continue working for this employer?"
Sample: Same as in Table 1.
Details: Means and standard errors are computed from multiple implicates, as in Rubin (1987), using survey weights. Years of potential experience is defined as age minus years of completed education minus six.

Approximately 14% of respondents answered that they would "never stop"; we imputed a specific answer for them as follows: (1) we used their answer if they responded to a later question about when they would retire from all work; or else (2) we used their answer if they responded to a later question about when they would retire from full-time work; or else (3) we assumed that they would work until the age of seventy.

In 1995-98, "less than a year" was coded as a separate answer, in which case we assigned a value of zero; in 1983-92 one is the smallest coded response, and for respondents who were coded with a value of one, we randomly assigned an answer of zero in the same proportion as is observed among those answering zero or one in 1995-98 (which will lead to a slight underestimate of the decline in tenure).

Table 3: Job tenure and pension coverage (OLS regression results, SCF)

	Dependent variable: years of current job tenure			
	Men (mean of dependent variable = 9.51)			
	(1)	(2)	(3)	(4)
Independent variables:				
has DB pension only	5.72*** (0.26)	4.35*** (0.28)	5.68*** (0.27)	4.54*** (0.28)
has DC pension only	2.56*** (0.25)	2.24*** (0.25)	2.57*** (0.25)	2.29*** (0.26)
has DB & DC pension	5.81*** (0.31)	5.05*** (0.32)	5.78*** (0.32)	5.06*** (0.32)
	Women (mean of dependent variable = 7.62)			
	(5)	(6)	(7)	(8)
Independent variables:				
has DB pension only	4.17*** (0.31)	3.47*** (0.32)	4.14*** (0.31)	3.54*** (0.32)
has DC pension only	2.24*** (0.25)	2.06*** (0.25)	2.23*** (0.25)	2.06 (0.24)
has DB & DC pension	4.46*** (0.35)	3.92*** (0.37)	4.47*** (0.35)	3.99*** (0.37)
Regression also includes:				
age	yes	yes	yes	yes
job variables	no	yes	no	yes
year effects	no	no	yes	yes
year*job variables	no	no	no	yes

Data source: Survey of Consumer Finances 1983, 89, 92, 95, 98, 01.
Sample: Full-time employees, excluding those who report tenure in excess of potential experience plus two (about 1.5% of the sample); those whose pension type is unknown (approximately 0.5% of the remaining sample); and those with earnings in the top or bottom 1% of the distribution.
Details: The coefficient estimates and Huber-White standard errors are computed from regressions run on multiple implicates, as in Rubin (1987). The regressions were weighted using survey weights. * indicates a confidence level of at least 90%, ** 95%, *** 99%.

Specifications: (1) and (5) include real weekly earnings (in 2001 dollars), age and age squared. (2) and (6) add job variables (4 education, 6 industry, 6 occupation, and 6 firm size dummies, industry* occupation, education*occupation, union coverage). (3) and (7) add year dummies. (4) and (8) add variables from (2) and (3) along with year*industry, year*occupation, year*education, year*union coverage.

Table 4: Job tenure and DB pension characteristics (OLS regression results, SCF)

	Dependent variable: years of current job tenure			
	Men (mean of dependent variable = 9.51)			
	(1a)	(4a)	(1b)	(4b)
Independent variables:				
has DB pension only	2.32*** (0.37)	1.11*** (0.46)	3.37*** (0.32)	2.71*** (0.45)
has DC pension only	2.68*** (0.25)	2.44*** (0.26)	2.64*** (0.25)	2.33*** (0.26)
has DB & DC pension	2.33*** (0.42)	1.55*** (0.50)	3.38*** (0.40)	3.16*** (0.51)
DB pension benefits at retirement (natural log of real present value, 1998 dollars):				
	Individual-reported		Average	
log value of monthly benefit	0.56*** (0.06)	0.57*** (0.07)	0.37*** (0.05)	0.28*** (0.07)
log value of lump-sum benefit	0.31*** (0.12)	0.35*** (0.12)	0.21* (0.12)	0.20* (0.12)
log value of pension wealth	0.55*** (0.05)	0.53*** (0.05)	0.43*** (0.05)	0.41*** (0.05)
log weekly earnings	1.33*** (0.23)	2.31*** (0.25)	1.38*** (0.23)	2.37*** (0.25)
	Women (mean of dependent variable = 7.62)			
	(5a)	(8a)	(5b)	(8b)
Independent variables:				
has DB pension only	1.05* (0.65)	0.23 (0.75)	1.82*** (0.51)	1.55*** (0.54)
has DC pension only	2.37*** (0.25)	2.25*** (0.24)	2.34*** (0.25)	2.17*** (0.24)
has DB & DC pension	1.30*** (0.60)	0.57 (0.71)	2.11*** (0.52)	1.95*** (0.55)
DB pension benefits at retirement (natural log of real present value, 2001 dollars):				
	Individual-reported		Average	
log value of monthly benefit	0.59*** (0.11)	0.64*** (0.14)	0.43*** (0.09)	0.38*** (0.10)
log value of lump-sum benefit	0.42*** (0.19)	0.46*** (0.20)	0.34** (0.19)	0.31** (0.19)
log value of pension wealth	0.36*** (0.08)	0.32*** (0.07)	0.22*** (0.07)	0.21*** (0.07)
log weekly earnings	1.69*** (0.29)	2.52*** (0.36)	1.80*** (0.30)	2.58*** (0.37)
Regression also includes:				
age	yes	yes	yes	yes
job variables	no	yes	no	yes
year effects	no	yes	no	yes
year*job variables	no	yes	no	yes

Details: These regressions replicate those appearing in Table 3, with the addition of variables representing DB pension benefits expected at retirement The value was reported in one of three different ways: (1) over 95% of individuals with a DB pension in 1989-01 reported a periodic amount that they expect to receive when they leave their job; (2) about 2.5% of individuals with a DB pension in 1989-01 reported a lump-sum amount which they expect to receive; (3) the SCF reported expected pension wealth for 55% of individuals with a DB pension in 1983, based on information collected from employers. We included the natural log of the present value of each of these variables, along with dummy variables indicating which of the three variables (if any) was reported for a given observation. In regressions (1a), (4a), (5a), and (8a), the self-reported variable is included. In regressions (1b), (4b), (5b), and (8b), the average value is included, imputed on the basis of log earnings, industry, occupation, education, unionization, and employer size, separately for men and women.

The numbering of the regressions parallels the numbering in Table 3. Huber-White standard errors appear in parentheses; * indicates a confidence level of at least 90%, ** 95%, *** 99%.

For additional information, see notes to Table 3.

Table 5: Changes in the value of DB pensions (OLS regression results, SCF)

Dependent variable: expected monthly pension benefit (2001 dollars)

	Men (mean of dependent variable = 1637)	
	(1)	(2)
Independent variables:		
year dummy, 1992	-32 (187)	-70 (201)
year dummy, 1995	-150 (161)	-253 (178)
year dummy, 1998	-222 (170)	-352* (187)
year dummy, 2001	-107 (193)	-202 (202)
	Women (mean of dependent variable = 1033)	
	(3)	(4)
Independent variables:		
year dummy, 1992	-333*** (134)	-357*** (145)
year dummy, 1995	-40 (218)	-82 (245)
year dummy, 1998	-266* (141)	-294** (130)
year dummy, 2001	-224 (162)	-246 (161)
Regression also includes:		
age, tenure, experience	yes	yes
job variables	no	yes

Data source: Survey of Consumer Finances 1989, 92, 95, 98, 01.
Sample: Full-time employees with DB pensions who report their expected monthly benefit, excluding those who report tenure in excess of potential experience plus two (about 1.5% of full-time employees); those whose pension type is unknown (approximately 0.5% of the remaining sample); those with earnings in the top or bottom 1% of the distribution; those who report that they will receive a lump-sum benefit (2.5% of the remaining sample) and those who do not report a benefit (2.5% of the remaining sample).
Details: The coefficient estimates and Huber-White standard errors are computed from regressions run on multiple implicates, as in Rubin (1987). The regressions were weighted using survey weights. * indicates a confidence level of at least 90%, ** 95%, *** 99%.

Specifications: (1) and (3) includes real weekly earnings (in 2001 dollars), age and age squared, potential experience and experience squared, current tenure (linear through quartic terms), and expected future tenure (linear through quartic terms). (2) and (4) add job variables (4 education, 6 industry, 6 occupation, and 6 firm size dummies, industry* occupation, education*occupation, union coverage).

Table 6: The earnings premium associated with job tenure

	Men						
	1983	1987	1991	1996	1998	2000	2002
Years of tenure							
5	13.9% ***	19.5 ***	15.0*** [87]	14.7*** [87]	10.4*** [91]	11.1*** [87]	8.8*** [96]
10	21.3 ***	27.6 ***	22.5*** [87]	23.5 ***	17.5*** [96]	16.5*** [96]	14.1*** [96]
15	24.6	30.7 ***	27.0 ***	28.6 ***	22.7*** [96]	20.4*** [96]	18.3*** [96]
20	25.9	33.1 ***	30.6 *	31.8 *	27.0 ** [87]	24.5 ** [96]	22.3 ** [96]
25	27.4	36.6	33.4	34.7	30.5 [87]	27.9 [96]	26.0 [96]
	Women						
Years of tenure							
5	17.5 ***	19.6 ***	19.3 ***	16.4 ***	9.1*** [96]	9.0*** [96]	5.3*** [96]
10	24.0 ***	29.4 ***	26.4 ***	25.3 ***	18.1*** [96]	14.9*** [96]	11.4*** [98]
15	27.9 ***	34.0	30.2 ***	30.5 ***	26.3*** [87]	18.7 ** [98]	16.8*** [98]
20	33.3 ***	36.2	35.1 ***	34.7 **	32.9 **	21.5 [98]	20.7 [98]
25	40.6	37.7	41.6	39.0	36.8	24.2 [98]	23.5 [98]

Data source: Outgoing rotation groups of the Current Population Survey tenure supplements of January 1983, 1987, 1991, February 1996, 1998, and 2000, and January 2002.
Sample: Employees aged 21-59 who were working or had a job but were not at work, excluding those who report earnings in the top or bottom 1% of the distribution. The sample size is 37,302 men and 36,024 women.
Details: Each cell in this table reports the estimated effect of years of job tenure on the natural log of the real wage, expressed as a percentage increase associated with a given number of years of tenure. These estimates are obtained from regressions run separately on men and women. The regressions include years of job tenure (with nonlinear terms included up to the fourth power), all interacted with the CPS year; years of potential experience (up to the fourth power), all interacted with the CPS year; a dummy for being a usual full-time worker; dummies for four education categories, all interacted with the CPS year; 51 industry dummies, interacted with a dummy for being in the public sector; and 45 occupation dummies, interacted with a dummy for being in the public sector. The real wage is defined as weekly earnings divided by usual weekly hours.

All of the estimated earnings premia are significantly different from zero. Two additional indications of statistical significance are reported; asterisks indicate significance of differences in the tenure premium across years of tenure in the same calendar year, while years indicate significance of differences in the tenure premium across calendar years for the same year of tenure. The asterisks next to each value indicate the significance level on a test that compares the earnings premium in that cell with the one reported in the cell below; * indicates a confidence level on the F-statistics of at least 90%, ** 95%, *** 99%. The years reported in brackets next to each value report the most recent year in which the value in the cell is significantly lower (at least at the 90% confidence level) than the value in the cell of a previous year. These estimates are obtained from regressions run separately on men and women, weighted using the outgoing rotation group weights.

Table 7: Technological change and long-term jobs (OLS regression results, April 1993 CPS)

Independent variables (averages, by industry):	Dependent variable (averages, by industry): % with a DB pension (1)	(2)	(3)	Average job tenure (4)	(5)	(6)
computer use, 1993	-0.095	-0.183	-0.216**	-4.00***	-3.99**	-4.72***
	(0.080)	(0.113)	(0.083)	(1.25)	(1.78)	(1.25)
past changes in computer use:						
1984-89	-	0.370	-	-	-0.09	-
		(0.332)			(5.23)	
1989-93	-	0.041	-	-	0.13	-
		(0.557)			(8.77)	
Regression also includes: firm size dummies	no	no	yes	no	no	yes
(averages over past 10 years, by industry)	(7)	(8)	(9)	(10)	(11)	(12)
TFP growth	-3.41***	-3.43***	-3.54**	20.7	19.7	7.0
	(1.22)	(1.23)	(1.31)	(22.7)	(22.9)	(27.0)
level of investment/10^6	-0.50**	-	-0.64*	-6.93*	-	-9.45
	(0.22)		(0.32)	(4.12)		(6.47)
capital stock/10^6	-	-0.24**	-	-	-3.46*	-
		(0.10)			(2.02)	
Regression also includes: firm size dummies	no	no	yes	no	no	yes

Data source: Current Population Survey, tenure, compensation, and firm size data from April 1993; computer use data from October 1984, 1989, 1993, and 1997. Jorgenson Total Factor Productivity Series, 1958-1996, obtained at http://post.economics.harvard.edu/faculty/jorgenson/data/35klem.html (Jorgenson, Gollop, and Fraumeni 1987).

Sample: The underlying samples from the CPS consist of workers aged 18-64 who are not self-employed. For the computer use data, sample sizes range from 49,601-54,647. For the April 1993 CPS data, the sample is restricted further to those who know their pension status, and the sample size is 12,951. Both data sets are then collapsed by taking averages within each of 50 industries, using the appropriate CPS sampling weight.

Details: Each numbered column reports coefficient estimates, standard errors in parentheses, and the significance level (* 90%, ** 95%, *** 99%) from an OLS regression where the dependent variable is defined at the top of the column, some of the independent variables are reported in the rows, and the other independent variables are log weekly wages, a dummy for whether the employer offers health insurance, and, where noted, dummies for the size of the employer (which was reported in this tenure supplement but not others used in Table 8). Each regression is weighted using the number of people in the industry in the April 1993 CPS. Capital and investment are deflated to 1992 dollars.

Table 8: Technological change and long-term jobs
(OLS regression results, multiple CPS tenure supplements)

Independent variables (averages, by industry):	Dependent variable (averages, by industry): Average job tenure (1)	(2)	(3)	(4)	(5)	(6)
computer use	-1.09**	-0.77	-1.04**	-0.67	-0.58	-1.10*
	(0.48)	(0.56)	(0.48)	(0.61)	(1.11)	(0.67)
past change in computer use	-	-	0.91	-0.80	-	-
			(1.14)	(1.89)		
Includes year dummies		X		X	X	X
Sample	1983, 87, 91, 96, 98, 00				<1990	>1990
(averages over past 10 years, by industry)	(7)	(8)	(9)	(10)	(11)	(12)
TFP growth	9.0	14.5	7.6	12.7	28.9**	-0.3
	(10.9)	(11.4)	(10.8)	(11.4)	(16.1)	(16.6)
level of investment/10^6	-8.6***	-7.4***	-	-	-4.8	-9.2***
	(2.6)	(2.7)			(4.7)	(3.4)
capital stock/10^6	-	-	-4.6***	-4.0***	-	-
			(1.3)	(1.3)		
Includes year dummies		X		X	X	X
Sample	1983, 87, 91, 96				<1990	>1990

Data source: Current Population Survey, tenure and compensation data from January 1983, 1987, and 1991 and February 1996, 1998, and 2000; computer use data from October 1984, 1989, 1993, and 1997. Jorgenson Total Factor Productivity Series, 1958-1996, obtained at http://post.economics.harvard.edu/faculty/jorgenson/data/35klem.html (Jorgenson, Gollop, and Fraumeni 1987).
Sample: The underlying samples from the CPS consist of workers aged 18-64 who are not self-employed. For the computer use data, sample sizes range from 49,601-54,647. For the tenure data, sample sizes range from 13,389-56,401. Both data sets are then collapsed by taking averages within each of 50 industries, using the appropriate CPS sampling weight.
Details: Each numbered column reports coefficient estimates, standard errors in parentheses, and the significance level (* 90%, ** 95%, *** 99%) from an OLS regression where the dependent variable is average job tenure, some of the independent variables are reported in the rows, and the other independent variable is log weekly wages. Computer use data in a given year are constructed as linear combinations of the averages from the computer use supplements immediately preceding and following that year, where the weights depend on the number of years. Each regression is weighted using the number of people in the industry in that year. Capital and investment are deflated to 1992 dollars.

Figure 1: Accrual of Pension Wealth

Figure 2

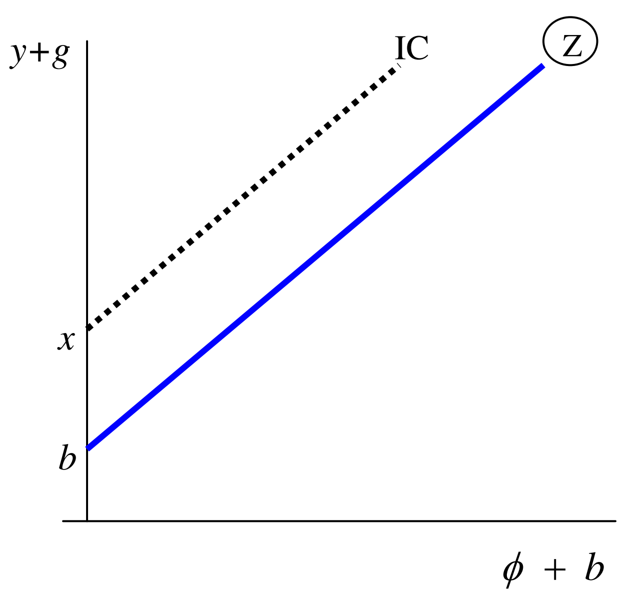

Figure 3

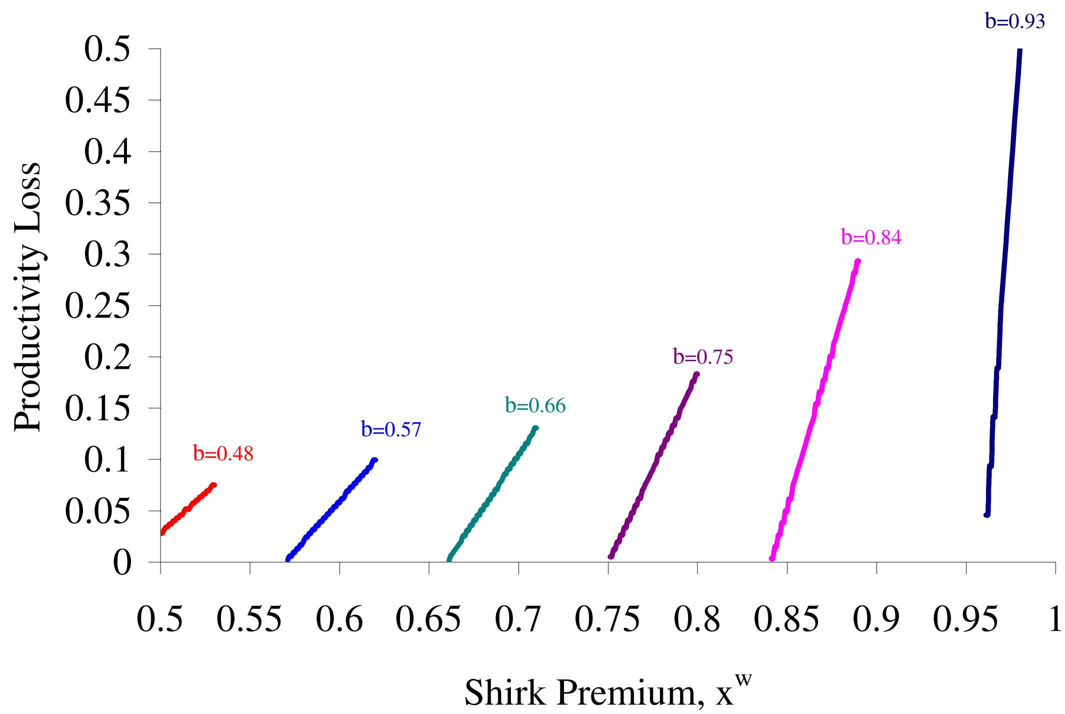

Figure 4