Earnings Instability and Tenure

Lorenzo Cappellari and Marco Leonardi*
Catholic University of Milan and University of Milan

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Abstract

This paper develops a tractable empirical approach to estimate the effect of on-the-job tenure on the permanent and the transitory variance of earnings in Italy. The model is also used to evaluate earnings instability associated with fixed-term contracts (short-tenure contracts). Our results indicate that each year of tenure on the job reduces earnings instability on average by 17%. Workers on a fixed-term contract can have an earnings instability up to 50% higher than workers on a permanent contract.

Keywords: Earnings instability, earnings dynamics, tenure, temporary contracts, minimum distance estimation.

JEL classification: J21, J31.

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1 Introduction

Estimating the changes in the variance of earnings is the topic of a vast body of research on earnings inequality and mobility. Many studies focus on cross-sectional evidence to investigate competing explanations of increasing inequality such as technical change, trade, institutions. A complementary strand of literature uses panel data on individual earnings to look at the extent of intertemporal mobility in the distribution of earnings, distinguishing long-term earnings components (e.g. ability) from temporary shocks. In this literature the individual earnings variance over time is typically studied as the sum of a permanent component (which has to do with changes in the quantity and prices of permanent individual characteristics) and a transitory component that captures the extent of earnings instability.

Gottschalk and Moffitt (1994) were the first to focus on the role of earnings instability in explaining inequality trends. Subsequent research extended Gottschalk and Moffitt's approach to several countries.[1] These studies found that, with varying intensities, earnings instability plays a role in explaining earnings inequality, and suggest that it may contribute to increase the extent of income uncertainty.

While the evolution of earnings instability in many countries is relatively well established, less is known about its determinants. Workers' mobility (job-to-job and in and out of unemployment) is a possible cause of earnings volatility, in particular when it leads to a decline in job security and job tenure. Some types of instability are the result of voluntary decisions by workers and families; for example, increased earnings instability among the young may reflect a productive search for better matches with firms that need their skills (Topel and Ward, 1992). On the other hand also involuntary displacement may affect earnings instability (Huff-Stevens, 2001).

Although there is some evidence of increased mobility from different sources, it has been difficult to establish the link between earnings instability and workers tenure because the empirical literature has found little evidence of a decline in the average tenure data in the U.S.[2] Differently from

[1] An incomplete list for the US includes Daly and Duncan (1997), Dynarski and Gruber (1997), Haider (2001) Shin and Solon (2008), Gottschalk and Moffitt (2009) and Moffitt and Gottschalk (2008). See also Dickens (2000), Ramos (2003) and Alessie and Kalwij (2007) on the UK, Baker and Solon (2003) on Canada, Cappellari (2004) on Italy.

[2] Violante (2002), Ljungqvist and Sargent (1998) and Bertola and Ichino (1995) describe the 1980s and 1990s as a period of increased economic "turbulence" characterized by a high rate of skill depreciation upon a job switch, which is consistent with growing earnings instability. Kambourov and Manovskii (2008a and 2008b) and Moscarini and Thomsson (2008) document an increase in occupational mobility in the U.S. during the '80s and

the U.S., in Italy and in many other continental European countries the diffusion of fixed-term contracts generates additional variance in tenure across cohorts and constitutes an "institutional" reason of shorter tenures. In this paper we use administrative panel data for Italy to document the decline of average tenure across cohorts and the diffusion of temporary contracts among younger cohorts of workers and to estimate the impact of on-the-job tenure on earnings instability.

A large literature in labour economics has focused on tenure as one of the determinants of wages. This literature has focused on estimating the average wage returns to tenure, trying to control for the endogeneity of tenure using different methods.[3] However, job mobility may affect not only the mean but also the instability of earnings and its effect may last for several periods after job change.

Our contribution to the earnings instability literature is the explicit modelling of tenure in models of earnings variance. We estimate the relationship of interest by exploiting variation in instability and tenure between and within narrowly defined birth cohorts over time. By doing so we provide a rich descriptive characterization of the relationship between earnings instability and tenure.

In addition to estimating the effect of tenure, as an alternative way of looking at the determinants of earnings instability we look at the impact of fixed-term contracts on instability. Fixed-term contracts are short-tenure contracts that typically last two or three years and can be renewed only once, which spread in many European countries in the Nineties. A large literature has studied the effect of fixed-term contracts on employment, job flows and wage levels but nobody has looked so far at their effects on earnings instability. Yet one of the main policy concerns about the diffusion of fixed-term contracts is their implications in terms of earnings instability and

'90s. Jaeger and Stevens (1999), Gottschalk and Moffitt (1999) and other contributions in the same *Journal of Labor Economics* special issue find little evidence of a decrease in workers' tenure. Farber (2008) finds however some decrease in tenure of older workers.

[3] The endogeneity of tenure is due to the fact that firm seniority is likely to be correlated with unobservable factors, such as the quality of the worker/firm match. Among the many studies, see Abraham and Farber (1987), Altonji and Shakotko (1987), Topel (1991), Topel and Ward (1992), Neal (1995), Lillard (1999), Altonji and Williams (2005), Dustmann and Meghir (2005). Parent (2002) exploits the different implications in terms of covariance structure of earnings to distinguish between human capital and matching theories. He finds that, in line with the human capital theory, those who start out with a lower wage in a job have a steeper tenure profile. While Parent (2002) looks at the relationship between tenure and long term earnings, in this paper we also assess the impact of tenure on the unstable component of earnings.

welfare because the temporary part of earnings variance is often uninsurable in presence of imperfect capital markets. In this paper we explicitly model the role of contract type (standard or fixed-term) in the decomposition of the wage variance and provide an estimate of the earnings instability associated with a fixed-term contract.

Our results indicate that workers with four years of tenure have on average an earnings instability almost 70% lower than workers with zero years of tenure or in other words each year of tenure on the job reduces earnings instability by approximately 17%. In recent years, in particular (but not only) young workers on fixed-term contracts have an earnings instability 50% higher than workers on permanent contracts.

Apart from their relevance from the policy point of view, these results are of interest in connection with various literatures. Other studies in the earnings instability literature have established that both displacement (Huff-Stevens, 2001) and voluntary job change (Hospido, 2009; Leonardi, 2004) may impact on the transitory variance of wages in the U.S. We cannot distinguish between voluntary and involuntary job moves but we are the first to provide an explicit way to model tenure in the analysis of earnings instability. While not directly comparable in terms of quantities, the results in this paper are in line with those obtained in structural models of job search (Jolivet, Postel-Vinay and Robin, 2006; Flabbi and Leonardi, 2008) which assume definite functional forms and agents' behaviour. Flabbi and Leonardi also find that an increase in mobility (in the job offer arrival rate while on the job) increases earnings cross-sectional variance in the U.S. although they do not find much of an impact on measures of lifetime inequality, implying that growing cross-sectional dispersion is driven by the unstable earnings component.

The rest of the paper proceeds as follows. Section 2 describes the institutional background. Section 3 describes the data with particular attention to the evolution of average tenure on the job and the diffusion of fixed-term contracts in Italy. Section 4 explains the statistical model. Section 5 describes the results and Section 6 concludes.

2 Institutional Background

Similarly to other European countries, labour market flexibility has increased in Italy over the last twenty years, through a series of measures which introduced various kinds of fixed-term and temporary contracts without changing the legislation on permanent, open-ended contracts. As a result the average tenure and its distribution in the population changed, particularly in younger cohorts more exposed to the new contracts. In Table 1 we document a different accumulation of tenure across cohorts. This potentially constitutes an interesting case study also for other countries although it does not constitute a natural experiment because the introduction of fixed term contracts was neither exogenous nor limited to specific sections of the population.

Temporary contracts are of various types and may have different implications for the covariance structure of earnings. Temporary contracts are typically used either as a buffer stock against downturns, as instruments of churning policies or as screening devices (especially apprenticeship or training contracts). Starting from this last type of contract, Italy is the Oecd country with the lowest probationary period and the second highest EPL index (OECD, 2004), the trial period ranges between one and eight weeks depending on the occupation. The shorter the period of probation, the larger the use of temporary contracts as screening devices.

In Italy the first wave of reforms of temporary contracts took place in the mid eighties with the introduction of apprenticeship or training contracts. This type of contract is widely used because it is convenient for employers for various reasons. Firstly, they have lower labour costs for apprentices and pay a wage that is set by national collective bargaining agreements at a level that is significantly lower than the norm. Also they pay social security contributions at a lower rate. Finally, firms pay no dismissal costs when contracts expire and this is why they are attracted to it as a useful substitute for other types of fixed-term contracts.[4]

Other types of fixed term contracts have always existed but the most important reform was the "Treu-Package" (named after the then minister of labour) which in 1997 legalised temporary work agencies and liberalised both apprenticeship and fixed-term contracts. Temporary work provided through agencies and fixed-term contracts are used as churning policies and buffer stock against downturns. They typically are more expensive than the standard open-ended contract but temporary workers provided by agencies

[4] The lower labour costs are intended to compensate firms for the training costs that they incur. However the training content of this type of employment is usually low, even if it is regulated by labour laws. Firms are required to share training costs by giving apprentices time off work (for a minimum number of paid hours) to attend external training courses that are provided by local authorities or accredited training institutes (and sponsored by the Regions) outside the premises of the firm. At the end of the training periods, each apprentice should receive a certificate for the qualification they have acquired in their field of work.

can be dismissed at will, while fixed term contracts have a legal duration of two years and can be renewed only once.

The upsurge in temporary contracts since 1997 decreased average tenure especially among young cohorts. However the implication with regards to wage covariances is not straightforward and may depend on the type of temporary contract. First, it is actually well known that temporary contracts might lower the commitment of employers to employees and vice-versa (Booth et al., 2002). It could be that employees never invest in any particular job because they know it is only temporary, and overall productivity suffers. Or it could be that with multiple chances, employees are able to find better matches than under the previous regime and productivity is enhanced overall. Wage covariances drop as employees change jobs more frequently. Then, even though there is more instability, the wage level is higher than it would be otherwise.

Second, the issue of selection into temporary contracts may be important. Worker heterogeneity enters into the decrease in tenure if for example the least stable workers are the first to be offered temporary contracts. In this case temporary contracts just act as a mechanism to sort workers into those who generally have short tenure and those who tend to stay at jobs longer. While overall this would have no impact on wage covariance, it would lower covariance for those on temporary contracts and raise it for those on permanent contracts. In this paper we do not address explicitly the selection of workers in temporary contracts for lack of convincing instruments, therefore the estimate of the difference in instability between permanent and temporary contracts may be considered an upper bound.

Finally, if fixed term contracts were used multiple times to substitute for open-ended contracts they could have no effect at all. The legal duration of a fixed term contract is of two years therefore the introduction or the increase of this type of contracts is bound to decrease tenure but if they are merely a re-labelling of an old arrangement, they could have no impact at all on wage covariances. This use of temporary contracts however should be limited because they can be renewed only once at maximum for a total duration of four years.

In this paper we provide a rich descriptive model of the effect of tenure an fixed term contracts on the wage covariance, however we do not establish the cause of earnings instability in the face of changes in institutional arrangements because we do not have exogenous variation in institutions. Thus we have to look at the results as descriptive and not proscriptive.

3 Data Description

The data are drawn from the Italian Social Security Administration (INPS) archives and span the years 1985-1999. The original dataset collects social security records of a 1/90 random sample of employees born on the 10th of March, June, September, and December of every year. The original archives only include information on private sector firms in the manufacturing and service sectors, therefore all workers in the public sector, agriculture and self-employment are excluded. This selection is common for administrative data which typically include the private sector only. Using the Bank of Italy data for 1998 (Survey of Households Income and Wealth, SHIW) the private sector constitutes 52% of total employment, agriculture represents only 2% while public employment and self employment represent 23% each. While there is evidence that wages are less volatile in the public sector compared to the private one (Cappellari, 2002), there are no studies on earnings instability among the self employed and agricultural workers, whose wages are likely to be more volatile than those of private sector employees.

The dataset contains individual longitudinal records generated using social security numbers. However, since the INPS collects information on private sector employees for the purpose of computing retirement benefits, employees are only followed through their employment spells in the private sector. The dataset stops following individuals who move into self-employment, the public sector, the agricultural sector, the underground economy, unemployment and retirement. In this paper we do not model selection from the private sector into other states (public sector, self-employment, unemployment and retirement) however the data on transition into other states say that workers are very stable in the private sector. After two years (always using SHIW data) 83% of employed male workers of age between 21 and 55 in 1998 are still working in the private sector, only 7.5% moved to the public sector, only 3% to self employment and 4.8% to unemployment and pension.

We have information on employees' age, gender, occupation (blue collar-white collar), yearly earnings, number of paid weeks, the initial and final month of job matches and the type of contract (permanent-temporary).

3.1 Sample selection rules

We keep in the sample all male workers aged 21 to 55 with positive yearly earnings and positive weeks of work. As customary in this literature, we focus on males since their labour force participation is less endogenously intermittent relative to females. The selection on age is aimed at avoiding the extremes of the working career, because employment volatility just after entry into the labour market or close to retirement may blur the measurement of structural earnings instability. In the course of the paper we use weekly earnings (yearly earnings divided by the number of weeks paid). For the cases of multiple individual spells in the same year we consider the longest spell.

The administrative data in electronic form start in January 1985 and the start date of all contracts already running at that date are artificially set to January 1985. Therefore, in order to measure tenure accurately, we consider only matches starting after the 1st of January 1985. Since such a selection rule leaves few observations in 1985 compared to the other years in the panel, we consider data from 1986 onwards. The final dataset includes 120,616 individuals with 632,105 person-year observations over the years 1986-1999.

The censoring of observations is necessary to calculate tenure and this is a common procedure with administrative data which do not contain information on tenure at entry (i.e. in the first year of the dataset). However about 21% of the observations and 9% of individuals are thrown out because of this (see Table 3), and it is possible that short tenure jobs (those with poor matches) will be over-represented in the sample because most stably employed men with long tenure in 1985 are likely to never show up in the data. Since employment generally becomes more stable at higher ages, more of the older men are likely to be totally excluded and this may lead to lower covariances. We approach this problem estimating a baseline model unconditional on tenure on both the censored and uncensored sample and checking the robustness of the results to sample selection.

Since our aim is to estimate tenure effects over a long period, it is crucial that in our models we control for age and time effects. To this end, we form subsamples defined by the year of birth and in order to ease the identification of tenure-related earnings profiles within each cohort we set the minimum period of observation of a cohort to five years. Given our sample selection on age, this implies that we consider cohorts of individuals born between 1935 (who turn 55 in 1990, in the fifth year of data in the sample) and 1974 (who turn 21 in 1995, and can be observed five times before the end of the sample). We therefore estimate the intertemporal covariance structure of earnings separately for these forty birth cohorts. Individuals born between 1944 and 1965 are observed fourteen times (i.e. over the whole sample period), while for individuals in cohorts born further apart from such interval the number of points in time monotonically decreases, going from 13 for those born in

1943 and 1966 to 5 for the oldest and youngest cohort, 1935 and 1974.

3.2 Descriptive statistics on tenure

Table 1 shows the average tenure in months in the full sample and within selected cohorts. All cohorts observed since the beginning of the panel start with low average tenure because the average refers only to contracts started after January 1985. Comparing the cohorts born in 1940-1960 from year 1986 up to year 1995 (i.e. the last year in which we observe the cohort born in 1940), it is evident that older cohorts accumulate on average longer tenure as a result of the lower job mobility of old workers relative to younger ones. In 1995, the difference between average tenure of the cohort born in 1940 and 1960 is 6.5 months. After 1995, the cohorts born in 1950 and 1960 remain in the sample, and their accumulation of tenure proceeds up until 1997, but stops in 1998 and 1999 (the years of the expansion of temporary contracts, the so called 'Treu Package'). The youngest cohort depicted in Table 1 (the cohort born in 1970) starts being observed in year 1991. Also in this case we can see that tenure accumulation is slower compared with older cohorts. For example, in 1997, after 7 years in the sample and before the reforms came into effect, this cohort accumulated 28.7 months of tenure, well below the comparable average tenure of the cohorts born in 1940, 1950 and 1960 in year 1993, i.e. after 7 years of observations. We exploit this variation of tenure across and within cohorts to estimate the model.

Table 2 shows the proportion of temporary workers and their average tenure for some of the youngest cohorts in our sample for whom the incidence of temporary contracts is more relevant. As explained in Section 2, the diffusion of temporary contracts is not limited to the late 1990s. In the late 1980s the so-called 'work and training' contracts (temporary contracts in which the employer had to pay reduced social security contributions and had to provide training on-the-job) were very popular, so that the overall share of temporary contracts reached 13 percent in 1988 for the cohort born in 1960. In 1997 the "Treu" reform introduced new forms of temporary employment, and our data show that their incidence increased substantially between 1997 and 1998 for the youngest cohorts. Compared to Table 1, Table 2 also shows that while permanent workers accumulate tenure on the job, the average tenure of temporary workers is always below 13 months. Moreover, their on-the-job tenure drops dramatically after the reform of the late 1990s, indicating that such reforms may have made temporary employment even more unstable than it was before.

Table 1: Average tenure in months

Year	Cohort born 1940	Cohort born 1950	Cohort born 1960	Cohort born 1970	Full sample	Obs.
1986	9.8	9.3	7.8		8.7	26413
1987	16.4	16.1	12.7		14.1	32170
1988	23.2	22.0	16.3		18.6	37361
1989	29.0	26.3	20.1		22.3	41120
1990	34.7	32.7	23.4		25.5	43030
1991	37.7	34.9	26.0	11.8	27.8	48322
1992	43.2	40.8	30.9	14.9	31.7	49588
1993	42.4	39.5	33.9	18.7	33.7	46795
1994	45.9	41.7	36.3	20.8	35.4	49254
1995	46.4	44.8	39.9	23.6	36.6	52233
1996		47.8	43.5	26.3	39.0	51897
1997		50.8	45.1	28.7	41.1	51553
1998		47.5	45.0	31.5	41.6	49681
1999		49.4	46.0	31.9	42.2	52688
N obs.						632105

Table 2: Incidence and average tenure in months of temporary contracts: Selected Cohorts.

Year	Cohort born 1960			Cohort born 1970			Cohort born 1974		
	obs.	%	Tenure temp.	obs.	%	Tenure temp.	obs.	%	Tenure temp.
1986	1056	0.07	6.4						
1987	1267	0.1	8.7						
1988	1449	0.13	9.5						
1989	1494	0.11	11.7						
1990	1491	0.04	17						
1991	1636	0.01	16.4	1772	0.2	8.7			
1992	1643	0.01	9.1	1892	0.19	11.7			
1993	1524	0.01	11.1	1807	0.16	12.8			
1994	1583	0.01	12.3	1916	0.13	10.9			
1995	1611	0.01	13.1	2019	0.12	10.8	1583	0.14	7.4
1996	1614	0.01	6.3	2106	0.11	10.7	1760	0.17	9.5
1997	1631	0.02	9.1	2168	0.12	11.1	1945	0.17	10.9
1998	1588	0.05	3.5	2102	0.16	5.2	1973	0.29	7.3
1999	1714	0.06	3.6	2397	0.17	5.2	2245	0.29	7.7

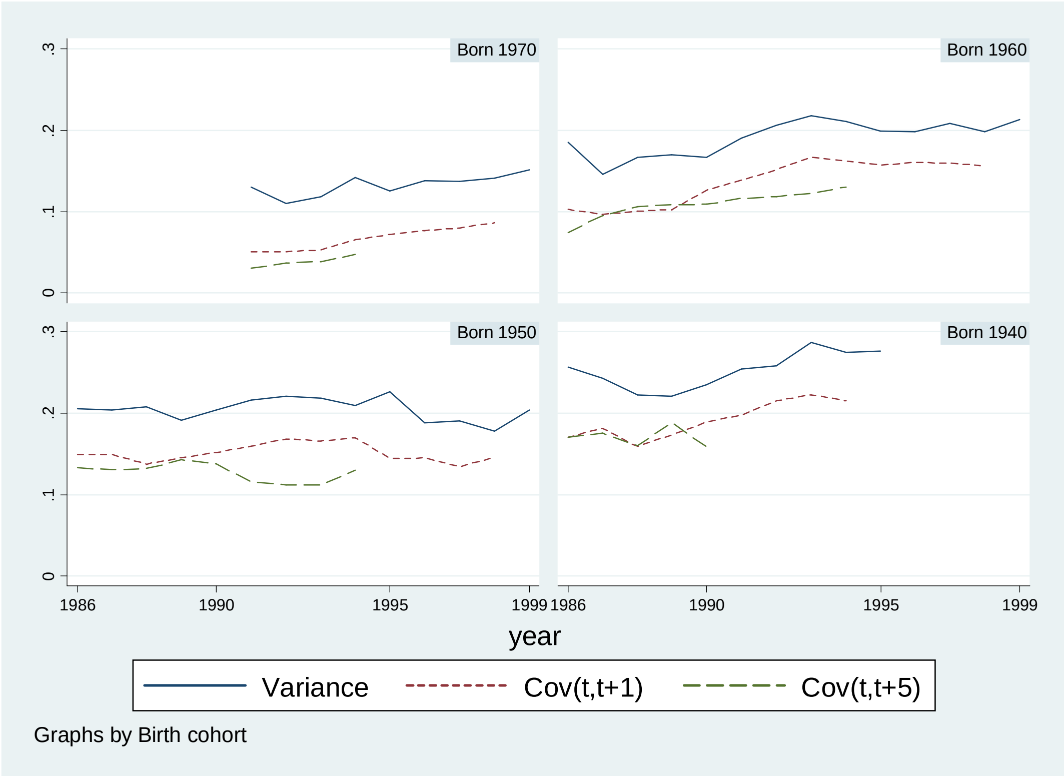

Figure 1: Raw moments

3.3 The intertemporal covariance structure of earnings

We use all valid wage observations in our sample to estimate the covariance structure of earnings for the forty birth cohorts. While not solving issues of endogenous panel attrition, such an unbalanced panel design is certainly less restrictive compared with analyses based on balanced panels.[5]

We plot estimated variances and covariances for selected birth cohorts in Figure 1. For most of the cohorts, earnings dispersion appears to increase at a moderate pace over the initial part of the period investigated, while trends of earnings inequality seem to level off towards the end of the sample period. These patterns reproduce the evidence for Italy provided by other

[5] Approximately $120,000$ individuals in our sample are observed over the whole period investigated. We estimate earnings second and fourth moments from the unbalanced panel following the procedure described in Dickens (2000).

studies, see e.g. Brandolini et al. (2002). Covariances at various lags are at a lower level compared with the variance, but still show the upward trends discussed above. The distance between covariances at increasing lags, moreover, decreases over lags, and covariances tend to stabilize to a long-term level. Such a pattern is consistent with an underlying process of earnings dynamics formed by some long-term component plus some mean-reverting component characterized by low order autoregression. We now turn to the modeling of such processes.

4 Econometric Model

We characterise the link between earnings instability and tenure by modelling the intertemporal covariance structure of earnings. We define earnings instability as the variance of the transitory component of earnings.

Our data enable us to observe forty sub-samples defined according to the year of birth over the period 1986-1999, yielding the possibility to separate time and birth cohort effects.[6] We achieve this by estimating the cohort-specific earnings covariance structure and jointly modelling the second earnings moments of all cohorts. In particular, let w_{ijt} be the deviation of log-earnings for individual i in job j and year t from the cohort- and period-specific mean, with $i = 1 \ldots N$ and $t = s_i, \ldots, S_i$. The initial and final point of observation, s and S, are individual specific because the presence in the panel depends upon birth cohort membership (see Section 3).

Earnings differentials within each cohort can be analysed by modelling the earnings covariance structure $E(w_{ijt}\ w_{ihk}), t - k = 0, ..K_i$. Given the unbalanced design by cohort, the longest time interval over which earnings covariances can be estimated is individual specific.

4.1 Benchmark model

We start by characterising the benchmark decomposition of earnings differentials between earnings instability and long-term persistence. As in Baker and Solon (2003), the base model allows for age effects in both permanent and transitory earnings, ensuring that the tenure effects in which we are interested in - introduced later on in this Section - will not pick up age-related earnings dynamics.

[6] Estimating age, time and cohort effects would require some parametric restrictions, see Alessie and Kalwji (2006).

Let w_{ijt} be the sum of two orthogonal components, the long-term one (w_{ijt}^P) and the mean-reverting shock (w_{ijt}^T):

$$w_{ijt} = w_{ijt}^P + w_{ijt}^T; \qquad E(w_{ijt}^P w_{ijt}^T) = 0 \tag{1}$$

where the first component represents those earnings determinants that depend on long-term personal attributes such as education or learning ability on-the-job; the second component captures in each year the deviations of individual earnings from the person-specific long-term component. The orthogonality assumption allows separate identification of the two components.

In this basic set-up, we allow long-term earnings to evolve over the life cycle according to a random walk process and to be shifted by a period-specific loading factor π_t that accounts for aggregate shifts in the long-term earnings distribution (π_{1986} is normalized to one for identification):

$$w_{ijt}^P = \pi_t r_{ia(t)}; \quad r_{ia(t)} = r_{ia(t-1)} + \xi_{ia(t)}; \quad r_{i21} \sim_i (0, \sigma_r^2); \quad \xi_{ia(t)} \sim_{ia} (0, \sigma_\xi^2) \tag{2}$$

where $a(t)$ denotes age at time t, σ_r^2 captures heterogeneity in long-term earnings at age 21 (the youngest age in our sample), while σ_ξ^2 measures the dispersions of innovations to the earnings process over the life-cycle.[7] The random walk specification allows for age effects in long term earnings. Specifically, given the initial condition of the process r_{i21}, subsequent evolution is determined by the arrival of random shocks, that induce a permanent relocation of individuals in the distribution of long-term earnings as the life-cycle evolves.

In principle, other specifications of long-term earnings may be used. In particular, a popular one in the literature on earnings dynamics is the random growth model, based on heterogeneous linear wage profiles in age or experience (Hause, 1980; Baker, 1997; Haider, 2001; see Baker and Solon, 2003, for a model that mixes random walk and random growth). However, that model predicts a convex evolution of long-term dispersion with age, a feature that is not present in our data. Therefore, we use a random walk specification, which implies a linear relationship between the permanent variance and age. To see this, iterating the recursion back to age 21, one can rewrite long-term earnings as:

$$w_{ijt}^P = \pi_t (r_{i21} + \sum_{z=22}^{a(t)} \xi_{iz}) \tag{3}$$

[7] Random walk models of long-term earnings have recently been used by Dickens (2000).

so that the resulting covariance structure of long term earnings is:

$$E(w^P_{ijt}w^P_{ihk}) = (\sigma^2_r + min(a(t) - 21, a(k) - 21)\sigma^2_\xi)\pi_t\pi_k \quad (4)$$

i.e. the model implies a linear growth of long-term earnings dispersion over the life cycle.

For the volatile component we assume a non-stationary AR(1)[8] process and we allow for non-stationarity modelling the (variance of the) initial conditions of the autoregressive process:

$$w^T_{ijt} = \tau_t\mu_c v_{it}; \qquad v_{it} = \rho v_{it-1}+\varepsilon_{it}; \qquad \varepsilon_{it} \sim_{it} (0;\sigma^2_{\varepsilon ct}); \qquad v_{is} \sim_i (0;\sigma^2_0) \quad (5)$$

where $c = c(i)$ is the cohort index for individual i. Period- and cohort-specific shifters τ_t and μ_c are allowed for in order to control for aggregate shifts in the distribution of transitory earnings.[9] As for long-term earnings, we allow for age effects also in earnings instability. Specifically we model the variance of AR(1) innovations as a function of aggregate age measures, namely:

$$\sigma^2_{\varepsilon ct} = \sigma^2_\varepsilon exp[g_1(A_{ct})] \quad (6)$$

where A_{ct} is the average age of birth cohort c in year t, while $g_1(.)$ is a linear spline with knots at 29, 37 and 45, with associated coefficients γ_{a1} to γ_{a4}. Thence, we exploit variation in average age across periods and cohorts to identify its impact on earnings instability. Note that time and cohort effects are already controlled for through the non-parametric shifters τ_t and μ_c, so that $g_1(.)$ will pick up variation in instability due to age, and not to time and cohorts effects.[10]

It follows that:

[8] We also experimented with ARMA(1,1) specifications. However, when we modelled the impact of tenure on instability, moving average components proved difficult to identify, possibly because much of the serial correlation in the volatile component was absorbed by the coefficients on tenure. For the sake of comparability, we therefore adopt the AR(1) specification throughout the paper. Baker and Solon (2003) report similar issues in a model of instability without tenure.

[9] While other authors have used cohort specific variance of initial conditions –see e.g. Haider (2001) and Baker and Solon (2003)– here we allow the overall process to shift with birth cohort.

[10] Baker and Solon (2003) parameterise earnings instability using a quartic in age, and exploiting variation in age across cohorts and time periods, as we do. Our exponential specification ensures non-negativity of the age-related component, while preserving flexibility through the spline function.

$$E(w^T_{ijt}w^T_{ihk}) = \{d_0\sigma_0^2 + d[\sigma^2_{\varepsilon ct} + E(\upsilon^2_{it-1})\rho^2] + (1-d_0-d)[E(\upsilon_{it-1}\upsilon_{ik})\rho]\}\tau_t\tau_k\mu_c^2 \tag{7}$$

where $d_0 = I(t = k = s)$, $d = I(t = k > s)$, s is the starting year in the panel and $I(.)$ is an indicator function.

The orthogonality assumption given in 1 implies that the theoretical covariance structure of this model results from the sum of 4 and 7.

4.2 Modelling the impact of tenure

Our specific interest is in the impact of on-the-job tenure on earnings variance components. In principle, both components may vary with tenure. Long-term earnings may vary with tenure due to match-specific shocks on productivity. The random walk in tenure may emerge in a matching model in which new information on match productivity gets released in each period, inducing a permanent re-shuffling in the distribution of long-term earnings. Similar predictions, in terms of increasing long-term dispersion with tenure on-the-job, may arise in human capital model in which the accumulation of firm-specific skills follows the unit root process.

Parent (2002) stresses that training model would predict time-varying match-specific effects, and that, within the match, the correlation between wage growth and initial earnings may depend on the presence of selection effects of better matches into training. Parent (2002) models time-varying match effects using a random growth specification; in this paper, instead, for the reasons discussed above we resort to a match-specific random walk model, which complements the random walk in age:

$$\begin{aligned} w^P_{ijt} &= \pi_t(r_{ia(t)} + q_{ijt}) \\ r_{ia(t)} &= r_{ia(t-1)} + \xi_{ia(t)}; \quad r_{i21} \sim_i (0, \sigma_r^2); \quad \xi_{ia(t)} \sim_{ia} (0, \sigma_\xi^2) \\ q_{ijt} &= q_{ijt-1} + \psi_{ijt}; \quad q_{iy} \sim_i (0, \sigma_q^2); \quad \psi_{ijt} \sim_{it} (0, \sigma_\psi^2) \end{aligned} \tag{8}$$

where y denotes the initial period of match j, σ_q^2 measures long-term earnings heterogeneity at the start of the match, and σ_ψ^2 measures the dispersion of innovations to long-term earnings occurring during the match. The specification in 7 implies that now long-term inequality evolves linearly also over match duration:

$$E(w^P_{ijt}w^P_{ihk}) = \pi_t\pi_k(E(r_{ia(t)}r_{ia(k)}) + E(I_i(j = h)q_{ijt}q_{ihk})) =$$
$$= \pi_t\pi_k[\sigma^2_r + min(a(t), a(k))\sigma^2_\xi + E(I_i(j = h))(\sigma^2_q + \min(t - y, k - y)\sigma^2_\psi)] \tag{9}$$

where $I_i(j = h)$ is a dummy for job-stayers.

There are reasons to believe that also earnings instability should vary with job tenure. Specifically, we should expect earnings instability to decrease with job duration if the quality of the match is initially measured with error, or if firms are more willing to insure earnings against volatile shocks the more they know match quality (see Lange, 2007, and Guiso et al., 2005). While Parent (2002) models tenure effects only in the long term-component, in this paper we adopt a broader approach and assess the impact of tenure also on the unstable component of earnings. Specifically, we augment the specification of AR(1) innovations with a spline term in average (within period-cohort cells) tenure $g_2(.)$, with knots at 1, 2 and 3 years of tenure:

$$\sigma^2_{\varepsilon ct} = \sigma^2_\varepsilon exp[g_1(A_{ct}) + g_2(T_{ct})] \tag{10}$$

where T_{ct} denotes the average tenure for cohort c in period t, and the additional spline coefficients go from γ_{t1} to γ_{t4}. Substituting 10 into 7 and adding the result to 9 yields the earnings covariance structure for the model with tenure.

4.3 Modelling the impact of fixed-term contracts

As discussed in Section 3, much of the variation in tenure comes from the diffusion of temporary contracts. An alternative way to measure the relevance of firm seniority for earnings instability is to look at the type of contract, open-ended or fixed-term. The underlying idea is that fixed-term contracts are associated with job turnover and do not favour the accumulation of seniority, so that if tenure reduces instability, then we should expect larger instability on fixed term contracts relative to long-term ones. Moreover, to the extent that ability is distributed more homogeneously among fixed term workers, say because training investments are less frequent, the distribution of long-term earnings should be more compressed compared with employees on permanent contracts.

Therefore, alternatively to our main model of tenure, we model the effects of contract types on earnings instability by letting the variance of innovations to the transitory component in the benchmark model to shift with the proportion of workers on fixed-term contracts observed in a given cohort over time. More specifically we assume that:

$$\sigma^2_{\varepsilon ct} = \sigma^2_{\varepsilon} exp[g_1(A_{ct}) + \eta F_{ct}] \tag{11}$$

where F_{ct} is the proportion of fixed term contracts in cohort c and period t. Given the discussion above, we expect η to be positive.

For long-term earnings, we take an approach similar to the one used for instability, and allow their covariance structure to be a function of the incidence of fixed term contracts over cohort-period cells:

$$E(w^P_{ijt} w^P_{ihk}) = [\sigma^2_r + \min(a(t) - 21, a(k) - 21)\sigma^2_{\xi}]\pi_t \pi_{(t-k)} exp(\lambda F_{ct}) \tag{12}$$

4.4 Estimation

Let $\Omega(\theta)$ be the auto-covariance function implied by the earnings models, a function of an unknown parameter vector θ. We estimate θ by Minimum Distance (see Chamberlain, 1984; Haider, 2001). This is an application of the Generalised Method of Moments: the inter-temporal auto-covariance function of earnings implied by the model is mapped into empirical second moments of the inter-temporal distribution of earnings $M = N^{-1} \sum_i M_i$, M_i being the individual contribution to M. Let $m_i = vech(M_i)$ and $\omega(\theta) = vech[\Omega(\theta)]$. The parameter vector is identified by the following set of moment restrictions:

$$E[m_i - \omega(\theta)] = 0 \tag{13}$$

Given that some elements of θ are cohort specific, we derive empirical moments from the within cohort earnings distribution and stack them over cohorts in estimation (see Baker and Solon, 2003). We work with unbalanced samples, so that an individual contributes to the empirical moments of his cohort only when he has valid earnings (see Dickens, 2000, and Haider, 2001). Overall, we can exploit 3180 empirical moments, collected in the vector m. A consistent estimate of θ can thus be obtained from the empirical counterpart of the moments restriction in (12), i.e. by minimising the distance between empirical and "theoretical" moments:

$$\theta = \arg\min[m - \omega(\theta)]'[m - \omega(\theta)] \tag{14}$$

This is the so called Equally Weighted Minimum Distance estimator (EWMD). Altonji and Segall (1996) showed that although not efficient,

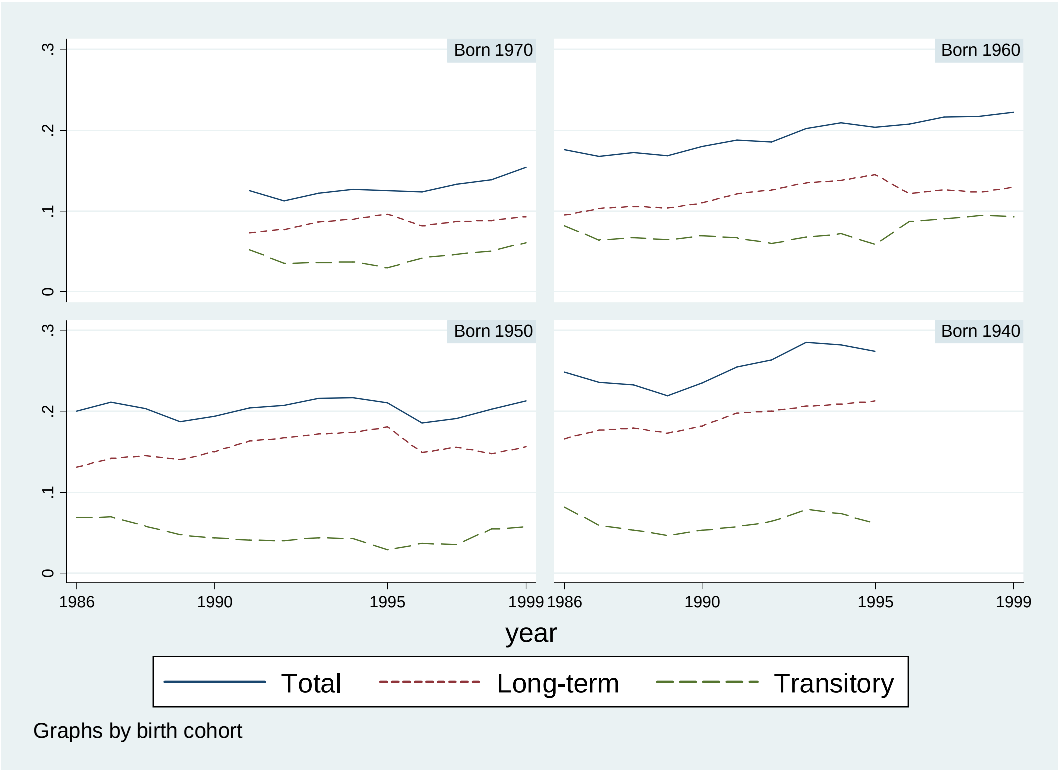

Figure 2: Predicted Variance Components

such estimator is preferable to the Optimum Minimum Distance (OMD, which uses $[var(m)]^{-1}$ to weight the minimisation problem) in the presence of correlations in sampling errors between second and fourth earnings moments. To reduce the variance of the estimator, we adjust standard errors using the fourth moments matrix after estimation, i.e. we use $var(\theta) = (G'G)^{-1}G'var(m)G(G'G)^{-1}$ to estimate the variance of estimated parameters, where G is the gradient of $\omega(\theta)$ evaluated at the solution θ^*.

In our tables of results we show a χ^2 statistic (with degrees of freedom equal to the number of moment conditions that exceed the number of model parameter) for the null hypothesis of correct model specification against the alternative of an unspecified covariance structure. Following Newey (1985), the statistic is obtained as $(m-\omega(\theta))'R^{-}((m-\omega(\theta))$ where R^{-} is the generalised inverse of $R = Wvar(m)W$, with $W = I - G(G'G)^{-1}G'$.

Table 3: Variance Components: Benchmark Model and Model with Tenure

	Censored sample				Non-censored sample	
	(1) Benchmark model		(2) Model with tenure		(3) Benchmark model	
	coeff.	s.e.	coeff.	s.e.	coeff.	s.e.
σ_r^2	0.0494	0.0017	0.0282	0.0033	0.0418	0.0011
σ_ξ^2	0.0042	0.0002	0.0035	0.0002	0.0036	0.0001
σ_q^2			0.0472	0.0084		
σ_ψ^2			0.0169	0.0027		
σ_ε^2	0.0673	0.0086	0.0728	0.0169	0.0492	0.0063
σ_0^2	0.1068	0.0119	0.0928	0.0130	0.0791	0.0089
ρ	0.3947	0.0112	0.2160	0.0263	0.4110	0.0090
γ_{a1}	0.0747	0.0101	0.0881	0.0242	0.0502	0.0089
γ_{a2}	0.0237	0.0077	0.0298	0.0118	0.0207	0.0068
γ_{a3}	-0.0618	0.0089	-0.0579	0.0123	-0.0615	0.0079
γ_{a4}	0.0369	0.0123	0.0699	0.0154	0.0414	0.0111
γ_{t1}			-0.5962	0.2648		
γ_{t2}			-0.5481	0.1462		
γ_{t3}			-0.0951	0.1766		
γ_{t4}			-0.8945	0.2352		
SSR	0.7168		0.6717		0.6827	
χ(d.f.)	8203 (3005)		7943 (2999)		10083 (3005)	
N indiv.	120,616		120,616		133,763	
N obs.	632,105		632,105		800,728	

5 Results

We begin our discussion by considering Figure 2 which plots for selected cohorts the variance decomposition into long-term and transitory components predicted by the model of equation 8. The predicted total variance of earnings replicates quite closely the patterns of the raw variance displayed in Figure 1, indicating that the fitting performance of the model is rather good. Earnings inequality increases as we move from younger to older cohorts, and long-term inequality seems to be the driving force behind such trends, whereas earnings instability remains fairly constant across cohorts.

These patterns suggest that we should expect age effects to be more evident in the long-term components of earnings inequality compared with the volatile one. The figure shows how these patterns of overall earnings inequality were driven by different factors over time. Increasing overall inequality in the late 1980s and early 1990s is essentially the result of widening long-term wage differentials, as would result from a widening distribution of skill premia, say in the presence of skill-biased technical change. Trends in the last part of the period analyzed have a different nature. While the level of permanent inequality drops between 1995 and 1996 and levels-off thereafter, earnings instability displays an upward pattern over the last years of observation, consistently with the increased labour market flexibility brought about by labour market reforms in this period.

5.1 Benchmark model

Parameter estimates of "core" earnings components are reported in Table 3, while for each model estimated period and cohort shifters are reported in the Appendix. We start our discussion of "core" parameter estimates by considering results for the benchmark model of Section 4.1, reported in column (1) of Table 3. The random walk coefficients (σ_r^2 and σ_ξ^2) measure heterogeneity in long-term earnings at the start of the working life and over the life-cycle and their estimates are statistically significant at any conventional level of confidence. In particular, our estimates imply that long-term dispersion almost quadruples between ages 21 and 55 ($(\widehat{\sigma}_r^2+\widehat{\sigma}_\xi^2*(55\text{-}21))/\widehat{\sigma}_r^2$ $=3.87$, a $\widehat{}$ denotes estimated coefficients), as a consequence of the heterogeneity in long-term earnings growth induced by the highly persistent shocks of the random walk model. Considering the earnings instability estimates for this benchmark model, one can observe a rather low degree of serial correlation in transitory shocks: the estimated AR(1) coefficient ρ implies that only 1 percent of a transitory innovation contributes to transitory earnings after 5 years. The dispersion of transitory shocks is substantial both at the beginning of the observation period (σ_0^2) and over time (σ_ε^2). The estimated coefficients of the exponential spline function in age (the γ_a's) are all precisely estimated, and reveal that the evolution of instability is not monotone over the life-cycle: instability grows at the beginning of the working life, decreases over its central part and then starts rising again towards the end of the life-cycle. A non-monotonic pattern of instability over age has also been found by Baker and Solon (2003) adopting a quartic specification.

5.2 Model with tenure

Parameter estimates for the main model of interest, laid out in equation 8, are reported in column (2) of Table 3. Considering the permanent component first, the additional (with respect to column (1)) coefficients reported (σ_q^2 and σ_ψ^2) summarise the evolution of long-term inequality with tenure. Job-specific long-term earnings inequality grows substantially within the firm-worker match. Our estimates imply that, considering someone starting a job at age 30, the level of long-term dispersion would increase by more than 75 percent during a job match that lasts four years ($\widehat{\sigma}_q^2$+$\widehat{\sigma}_\psi^2$*4+$\widehat{\sigma}_r^2$+$\widehat{\sigma}_\xi^2$*(34-21))/($\widehat{\sigma}_q^2$+$\widehat{\sigma}_r^2$+$\widehat{\sigma}_\xi^2$*(30-21)) =1.76).[11] This scenario is consistent with the arrival of new information on match productivity over time or the accumulation of firm-specific human capital, which widens the distribution of long-term earnings as matches elapse.

To the extent that tenure-related parameters capture within-match growth of permanent earnings inequality, age-related ones can be interpreted –in a residual sense– as picking up the changes in long-term inequality between jobs. Contrasted with their counterparts in column (1), the coefficients of the random walk in age are smaller, as a consequence of the fact that within-job growth of earnings heterogeneity is now captured by the tenure-related parameters. Still, our estimates point toward the existence of substantial age-related heterogeneity, which could result from the accumulation of general human capital. Considering an individual with four years of tenure at the end of the working life, parameter estimates still imply a more than triple increase in long-term earnings inequality between ages 21 and 55 ($\widehat{\sigma}_q^2$+$\widehat{\sigma}_\psi^2$*4+$\widehat{\sigma}_r^2$+$\widehat{\sigma}_\xi^2$*(55-21))/($\widehat{\sigma}_q^2$+$\widehat{\sigma}_r^2$) =3.47).

Parameter estimates for the transitory component confirm the evidence emerged from column (1) in terms of the AR(1) and age-related coefficients.

[11] Recall that we have a sample of new spells and that in our model instability depends on the average tenure within cohort-period cells, which is always shorter than four years, see Table 1.

The new parameters in column (2), the γ_t's, relate earnings instability with tenure, showing that instability decreases with seniority in the job. Specifically tenure decreases over the first two years of the worker-firm match, flattens out over the third year, and then starts decreasing again, and at a faster rate, over the fourth year. Importantly, the tenure effects that we estimate are obtained while controlling for the relationship between earnings instability and age (through the γ_a's) so that the result is net of any spurious influence that may emerge in the presence of correlation between age and tenure.

As for the permanent component we can compute the evolution of earnings instability after four years of tenure for an individual starting a job at age 30 and accumulating four years of seniority in that job. However, providing a closed form expression for such calculation is cumbersome in this case due to the recursive structure of the autocovariance function of transitory earnings, see equation 7. Therefore, for illustrative purposes, here we use parameter estimates to compute the variation with tenure of the variance of AR(1) innovations (i.e. the parameter that maps tenure effects into earnings instability), while leaving the computation of the impact of tenure on earnings instability for the discussion below. After some algebra, the estimated proportional change in $\sigma^2_{\varepsilon ct}$ after four years of tenure is $1 - exp(\widehat{\gamma}_{a2} * 4 + \widehat{\gamma}_{t4}) = -0.54$.

These effects indicate that earnings profiles stabilise as individuals settle down in their new jobs, and may be interpreted in a matching model framework in which earnings profiles tend to their long-term component as the quality of the match is revealed to employers (Gibbons and Farber, 1996; Altonji and Pierret, 2001).

Predictions from this model in terms of variance decomposition are summarised in Figure 3, which plots the core variance components, i.e. excluding time and cohort shifters, against age and tenure. Predictions are averaged over cohorts. The long-term component grows following a linear trend with both age and tenure, as predicted by the random walk specification, with some deviations from exact linearity that are due to the averaging over cohorts. The patterns of earnings instability, instead, are rather different. While the evolution with age seems to be constant over the life cycle, there is a clear downward trend with tenure. More specifically, the average instability is 0.091 at the start of the job match and 0.029 at the end of it, implying an yearly reduction rate of approximately 17%.

Results discussed so far have been obtained using a sample that excludes job spells starting before January 1st 1985. As discussed in the data section,

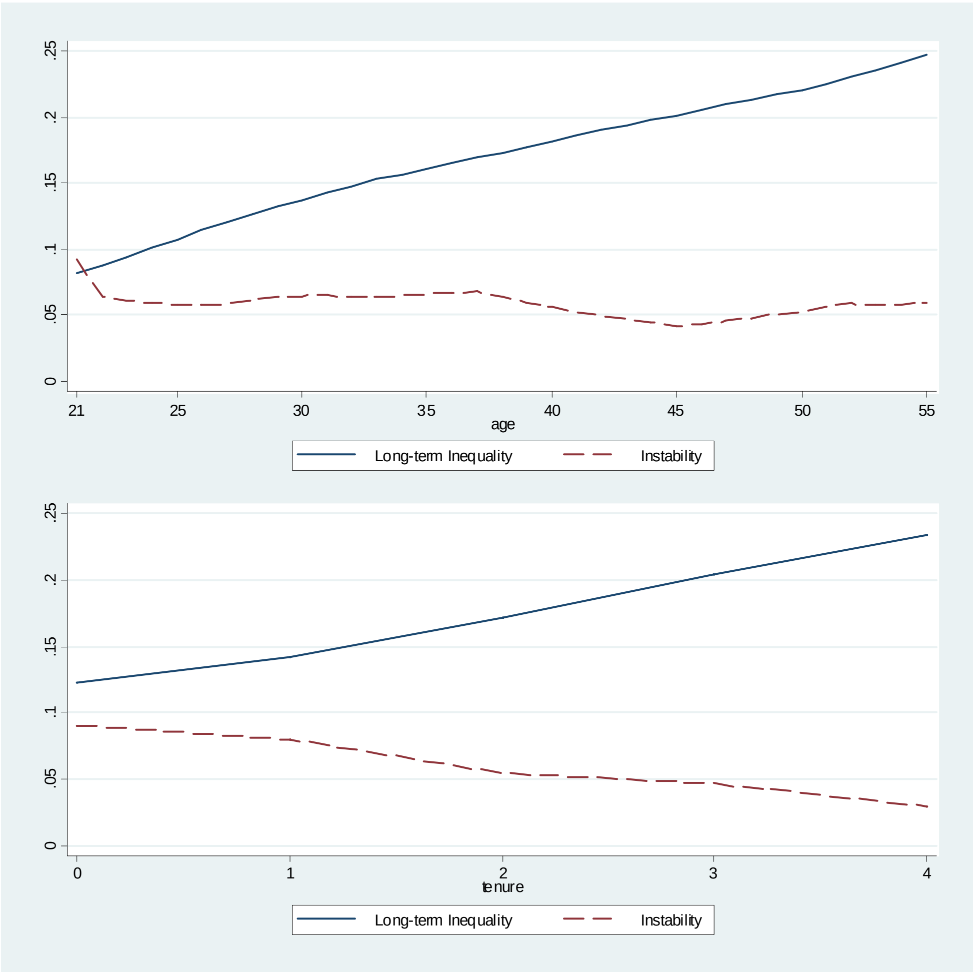

Figure 3: Long-term earnings dispersion and earnings instability, by age and tenure

a way to assess the robustness of our estimates to this type of sample selection is to re-estimate our model using a larger sample that also included cases with censored tenure. Necessarily, we run the robustness check in terms of the benchmark model, i.e. the one that does not use the tenure variable. Results are in column (3) of Table 3 and, overall, indicate that excluding censored spells does not seem to affect our estimates. Estimated parameters capturing heterogeneity (the variances of random walk in age; the variance of AR(1) initial conditions and the base variance of AR(1) innovations) are smaller in the non-censored sample, but quantitative differences are negligible in practical terms.

5.3 Model with fixed-term contracts

In the last part of Section 4 we discussed an alternative way to test our idea that shorter tenure is associated with earnings instability, namely to parameterise the variance components models with respect to the type of job contracts, fixed-term or open-ended. We do this by letting the dispersion of permanent and transitory earnings in the benchmark model to shift with the average proportion of temporary workers across period-cohort cells. Since time and cohort effects are already controlled for in the model by means of flexible loading factors, we are confident that the estimates will capture the association between variance components and contract type and will not be affected by other unobserved factors that vary by cohort and time period.

Results from this exercise are in Table 4, using both the censored and non-censored sample. Parameters estimates for coefficients other than the ones linking variance components to contract types are pretty similar to the ones estimated for the benchmark model, either on the censored and non-censored samples, and we do not comment them further. The coefficients linking contract type to permanent and transitory earnings shocks (λ and η respectively) attract the signs we would expect a priori, indicating that individuals on fixed term contracts have on average a lower permanent variance of earnings and a higher instability relative to permanent workers. The lower permanent variance probably reflects the lower heterogeneity of workers on fixed-term contracts in terms of age, education and of all observed and unobserved permanent characteristics. The higher transitory variance picks up the effect of the lower tenure (among other factors which are associated with a fixed-term contract and affect temporarily the wage).

Using parameter estimates we can compute the variation in earnings instability associated with fixed-term contracts (see Table 5). In particular, we consider the variation in instability that occurs when the proportion of

Table 4: Variance Components in Models of Fixed-Term Contracts.

	Censored sample		Non-censored sample	
	coeff.	s.e.	coeff.	s.e.
σ_r^2	0.0455	0.0015	0.0386	0.0010
σ_ξ^2	0.0030	0.0002	0.0027	0.0002
λ	-0.0148	0.0020	-0.0128	0.0016
σ_ε^2	0.0762	0.0089	0.0548	0.0063
σ_0^2	0.1142	0.0114	0.0850	0.0084
ρ	0.4472	0.0119	0.4581	0.0098
γ_{a1}	0.1049	0.0107	0.0772	0.0095
γ_{a2}	0.0387	0.0081	0.0327	0.0071
γ_{a3}	-0.0450	0.0078	-0.0470	0.0069
γ_{a4}	0.0378	0.0101	0.0419	0.0091
η	0.0371	0.0065	0.0327	0.0060
SSR	0.4754		0.4538	
$\chi(3003)$	9067		11580	
N indiv	120,616		133,763	
N obs	632,105		800,728	

fixed-term contracts changes from zero to the actual incidence observed in the data, net of time and cohort effects. We consider two periods with different diffusion of temporary contracts. In the year 1988, while for older cohorts born in 1940-1950 (where the actual incidence of fixed-term contracts is low) the resulting increase in instability is negligible, for younger cohorts earnings instability increases by almost 50 percent. In the most recent period (year 1998) characterized by a wider diffusion of temporary contracts also among older workers, the increase of instability is around 50% for the cohort born 1950 and proportionally larger for younger cohorts.

6 Discussion and Conclusion

In this paper we have used Italian panel data to estimate the impact of on-the-job tenure on earnings instability. Although other papers (Huff-Stevens,

Table 5: Predicted Earnings Instability by Contract Type.

	Cohort born 1940	Cohort born 1950	Cohort born 1960	Cohort born 1970
Year 1988				
permanent contract	0.078	0.089	0.070	-
temporary contract	0.081	0.094	0.106	-
Year 1998				
permanent contract	-	0.058	0.082	0.052
temporary contract	-	0.086	0.122	0.081

2001, Hospido, 2009 and Leonardi, 2004) have looked at the effect of voluntary and involuntary job changes on instability, we are the first to develop a formal model which accounts for tenure in the literature on the decomposition of earnings variance. We found that the dispersion of long-term earnings profiles increases with tenure while earnings instability declines with tenure. We estimate that each year of tenure is associated with a 17% reduction in instability. We also looked explicitly at the effect of fixed-term (short-tenure contracts) and permanent contracts on earnings instability. We found that workers on fixed-term contracts can experience between 10% and 50% more instability than the workers on permanent contracts.

These results are important from the policy point of view but are certainly relevant for various issues of general interest. In a different strand of literature which uses structural models of the labor market to establish a relationship between job mobility and earnings instability, Flabbi and Leonardi (2008) conclude that a large part of the increase in the transitory variance in the U.S. is correlated with the change in on-the-job arrival rate of job offers. While the results are not comparable from the quantitative point of view, the evidence in this paper goes in favor of the interpretation that changes in individual mobility are fundamental in explaining the evolution of wage dispersion.

The results in this paper are also consistent with different models of wage determination although we do not take a stance on the underlying theory. The results are consistent with matching models of wage determination where overall earnings profiles tend to their long-term component as individuals settle down in their job and information on their ability is revelead. A recent paper by Lange (2007) on U.S. data finds that the initial expectation error about match quality declines by 50% in three years which approximately equals our estimate of a reduction of 17% in earnings instability per year of tenure. Models of firm-provided insurance can also potentially account for these findings. Guiso et al. (2005) compute permanent and transitory shocks to firms' profits and workers' wages and find that firms provide workers with full insurance only against transitory shocks. This implicit contracts setting is consistent with our results if insurance provision grows with tenure and leads to a decline of earnings instability.

The exercise of this paper is particularly relevant for Italy, which starting from the late 1990s experienced an increasing diffusion of short term contracts. Many authors have stressed that the welfare effects of these reforms depend on their impact on employment probability. Here we have provided evidence that, even conditional on being employed, there may be additional channels through which these new type of jobs affect individual welfare, namely through an increased uncertainty surrounding long-term earnings profiles.

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Table 1: Appendix. Benchmark Model and Model with Tenure: time shifters 1986=1.

	Censored sample				Non-censored sample	
	Benchmark model		Model with tenure		Benchmark model	
	coeff.	s.e.	coeff.	s.e.	coeff.	s.e.
π_{1987}	0.9928	0.0109	1.0007	0.0115	1.0243	0.0065
π_{1988}	0.9661	0.0123	0.9905	0.0130	1.0287	0.0076
π_{1989}	0.9210	0.0127	0.9608	0.0130	1.0076	0.0078
π_{1990}	0.9196	0.0138	0.9782	0.0131	1.0266	0.0081
π_{1991}	0.9412	0.0147	1.0055	0.0142	1.0599	0.0088
π_{1992}	0.9241	0.0167	1.0155	0.0150	1.0719	0.0095
π_{1993}	0.9330	0.0189	1.0492	0.0165	1.0912	0.0113
π_{1994}	0.9230	0.0192	1.0481	0.0170	1.0955	0.0118
π_{1995}	0.9230	0.0197	1.0509	0.0173	1.1144	0.0122
π_{1996}	0.8231	0.0189	0.9358	0.0168	1.0060	0.0122
π_{1997}	0.8273	0.0195	0.9425	0.0173	1.0136	0.0130
π_{1998}	0.8080	0.0204	0.9236	0.0183	0.9891	0.0144
π_{1999}	0.8188	0.0205	0.9280	0.0188	1.0028	0.0149
τ_{1987}	0.8667	0.0758	0.8593	0.0352	0.9359	0.0316
τ_{1988}	0.9245	0.0752	0.8257	0.0374	0.9206	0.0358
τ_{1989}	0.9516	0.0835	0.7811	0.0366	0.8791	0.0358
τ_{1990}	1.0496	0.1053	0.7915	0.0380	0.9058	0.0382
τ_{1991}	1.0712	0.1208	0.7958	0.0397	0.9229	0.0402
τ_{1992}	1.1143	0.1511	0.7794	0.0398	0.9151	0.0409
τ_{1993}	1.1924	0.1754	0.7744	0.0410	0.9319	0.0430
τ_{1994}	1.2216	0.1972	0.7640	0.0420	0.9214	0.0440
τ_{1995}	1.1078	0.1964	0.7013	0.0414	0.8608	0.0435
τ_{1996}	1.3451	0.2593	0.7662	0.0456	0.9257	0.0479
τ_{1997}	1.4337	0.2972	0.7672	0.0483	0.9305	0.0505
τ_{1998}	1.5164	0.3203	0.7893	0.0510	0.9686	0.0540
τ_{1999}	1.5989	0.3453	0.8177	0.0542	0.9996	0.0569

Table 2: Appendix. Benchmark Model and Model of Tenure: cohort shifters 1952=1.

	Censored sample				Non-censored sample	
	Benchmark model		Model with tenure		Benchmark model	
	coeff.	s.e.	coeff.	s.e.	coeff.	s.e.
μ_{1935}	0.8307	0.1514	0.8572	0.1391	0.8367	0.1203
μ_{1936}	0.6218	0.1525	0.6644	0.1341	0.5744	0.1384
μ_{1937}	0.7932	0.1361	0.7993	0.1216	0.8127	0.1285
μ_{1938}	0.7447	0.1248	0.7619	0.1122	0.6829	0.1181
μ_{1939}	0.6579	0.1276	0.6857	0.1173	0.6562	0.1142
μ_{1940}	0.9399	0.1238	0.9404	0.1080	0.8825	0.1052
μ_{1941}	0.7724	0.1220	0.7820	0.1111	0.7634	0.1056
μ_{1942}	1.1321	0.1173	1.0816	0.1030	1.0872	0.0987
μ_{1943}	0.8963	0.1111	0.9036	0.1006	0.9458	0.0978
μ_{1944}	0.8532	0.1080	0.8356	0.0942	0.8304	0.0948
μ_{1945}	0.9201	0.1033	0.9266	0.0903	0.8798	0.0925
μ_{1946}	0.9061	0.1024	0.9188	0.0866	0.8796	0.0869
μ_{1947}	0.8908	0.1043	0.8974	0.0857	0.8843	0.0869
μ_{1948}	0.9255	0.1026	0.9358	0.0867	0.8568	0.0870
μ_{1949}	0.7796	0.1019	0.7957	0.0811	0.7787	0.0826
μ_{1950}	0.8617	0.0937	0.8933	0.0796	0.8613	0.0781
μ_{1951}	0.9625	0.0905	0.9626	0.0741	1.0081	0.0801
μ_{1953}	0.8923	0.0847	0.9249	0.0702	0.9753	0.0740
μ_{1954}	0.9431	0.0867	0.9709	0.0733	0.9966	0.0754
μ_{1955}	1.0109	0.0847	1.0411	0.0717	1.0744	0.0758
μ_{1956}	1.0443	0.0829	1.0571	0.0690	1.0628	0.0721
μ_{1957}	0.9157	0.0788	0.9739	0.0670	1.0081	0.0721
μ_{1958}	0.9375	0.0791	1.0103	0.0677	1.0327	0.0702
μ_{1959}	0.8583	0.0756	0.9566	0.0646	1.0067	0.0679
μ_{1960}	0.9358	0.0807	1.0360	0.0696	1.0659	0.0731
μ_{1961}	0.8340	0.0736	0.9361	0.0620	0.9903	0.0668
μ_{1962}	0.8152	0.0730	0.9349	0.0614	0.9880	0.0653
μ_{1963}	0.7857	0.0743	0.9131	0.0617	0.9565	0.0646
μ_{1964}	0.7538	0.0756	0.9018	0.0618	0.9579	0.0654
μ_{1965}	0.7105	0.0772	0.8838	0.0612	0.9646	0.0659
μ_{1966}	0.6922	0.0790	0.8975	0.0657	0.9469	0.0677
μ_{1967}	0.6710	0.0809	0.8897	0.0673	0.9424	0.0692
μ_{1968}	0.6598	0.0831	0.8795	0.0667	0.9291	0.0681
μ_{1969}	0.6749	0.0915	0.8998	0.0705	0.9406	0.0710
μ_{1970}	0.6999	0.1003	0.9433	0.0754	0.9739	0.0746
μ_{1971}	0.7297	0.1114	1.0051	0.0813	1.0214	0.0791
μ_{1972}	0.7171	0.1190	1.0424	0.0877	1.0502	0.0843
μ_{1973}	0.7179	0.1287	1.0846	0.0929	1.0795	0.0881
μ_{1974}	0.6864	0.1305	1.0669	0.0936	1.0455	0.0869

Table 3: Model with temp contracts: coefficients and standard errors of time shifters

	Censored sample		Non-censored sample	
	coeff.	s.e.	coeff.	s.e.
π_{1987}	0.9983	0.0116	1.0234	0.0066
π_{1988}	0.9907	0.0132	1.0289	0.0078
π_{1989}	0.9658	0.0134	1.0134	0.0081
π_{1990}	0.9849	0.0136	1.0322	0.0083
π_{1991}	1.0103	0.0146	1.0651	0.0090
π_{1992}	1.0198	0.0155	1.0762	0.0097
π_{1993}	1.0600	0.0172	1.1023	0.0119
π_{1994}	1.0662	0.0182	1.1127	0.0127
π_{1995}	1.0826	0.0194	1.1434	0.0138
π_{1996}	0.9628	0.0187	1.0311	0.0138
π_{1997}	0.9685	0.0192	1.0387	0.0146
π_{1998}	0.9460	0.0201	1.0105	0.0159
π_{1999}	0.9532	0.0210	1.0269	0.0167
τ_{1987}	0.8102	0.0320	0.9036	0.0289
τ_{1988}	0.7607	0.0337	0.8735	0.0328
τ_{1989}	0.7209	0.0327	0.8255	0.0327
τ_{1990}	0.7285	0.0338	0.8494	0.0344
τ_{1991}	0.7389	0.0354	0.8748	0.0363
τ_{1992}	0.7259	0.0356	0.8698	0.0370
τ_{1993}	0.7185	0.0370	0.8797	0.0390
τ_{1994}	0.6989	0.0376	0.8604	0.0399
τ_{1995}	0.6342	0.0371	0.7975	0.0397
τ_{1996}	0.6710	0.0406	0.8373	0.0432
τ_{1997}	0.6708	0.0428	0.8392	0.0455
τ_{1998}	0.6846	0.0453	0.8659	0.0490
τ_{1999}	0.6991	0.0480	0.8838	0.0517

Table 4: Model with termp contracts: coefficients and standard errors of cohort shifters

	Censored sample		Non-censored sample	
	coeff.	s.e.	coeff.	s.e.
μ_{1935}	0.9053	0.1252	0.8802	0.1051
μ_{1936}	0.7125	0.1246	0.6296	0.1246
μ_{1937}	0.8191	0.1087	0.8551	0.1136
μ_{1938}	0.8160	0.0966	0.7450	0.1001
μ_{1939}	0.7409	0.0971	0.7203	0.0937
μ_{1940}	0.9611	0.0956	0.9156	0.0923
μ_{1941}	0.8061	0.0960	0.7922	0.0904
μ_{1942}	1.0983	0.0935	1.0858	0.0879
μ_{1943}	0.9358	0.0829	0.9676	0.0824
μ_{1944}	0.8716	0.0818	0.8588	0.0806
μ_{1945}	0.9355	0.0780	0.8960	0.0779
μ_{1946}	0.9154	0.0748	0.8878	0.0749
μ_{1947}	0.9059	0.0749	0.8961	0.0754
μ_{1948}	0.9261	0.0745	0.8666	0.0745
μ_{1949}	0.8135	0.0715	0.8019	0.0721
μ_{1950}	0.8983	0.0719	0.8698	0.0702
μ_{1951}	0.9671	0.0681	1.0081	0.0726
μ_{1953}	0.9242	0.0639	0.9696	0.0662
μ_{1954}	0.9625	0.0656	0.9869	0.0667
μ_{1955}	1.0302	0.0646	1.0560	0.0667
μ_{1956}	1.0459	0.0617	1.0501	0.0636
μ_{1957}	0.9720	0.0609	0.9996	0.0644
μ_{1958}	1.0034	0.0610	1.0187	0.0619
μ_{1959}	0.9491	0.0578	0.9924	0.0596
μ_{1960}	1.0204	0.0621	1.0443	0.0639
μ_{1961}	0.9307	0.0555	0.9769	0.0587
μ_{1962}	0.9284	0.0550	0.9717	0.0572
μ_{1963}	0.9104	0.0556	0.9441	0.0569
μ_{1964}	0.9030	0.0561	0.9471	0.0580
μ_{1965}	0.8868	0.0558	0.9522	0.0584
μ_{1966}	0.9064	0.0606	0.9410	0.0607
μ_{1967}	0.9015	0.0625	0.9386	0.0624
μ_{1968}	0.8931	0.0623	0.9273	0.0616
μ_{1969}	0.9135	0.0664	0.9389	0.0648
μ_{1970}	0.9614	0.0717	0.9754	0.0686
μ_{1971}	1.0292	0.0784	1.0274	0.0736
μ_{1972}	1.0741	0.0859	1.0626	0.0797
μ_{1973}	1.1258	0.0923	1.0991	0.0844
μ_{1974}	1.1141	0.0943	1.0703	0.0842