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Earnings Instability and Tenure*

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Abstract

We study the effect of tenure on earnings instability in Italy using the reforms of temporary employment contracts, which affected the average tenure of workers differentially across cohorts. We develop a model of earnings dynamics, and we exploit the variation of tenure and instability over time and across birth cohorts to estimate policy-relevant parameters. Our results indicate that each year of tenure on the job reduces earnings instability by 11 percent; the drop is faster in the first three years of the match. Workers on a temporary contract have an earnings instability up to 100 percent higher than workers on a permanent contract.

Keywords: Earnings dynamics; labor reforms; permanent and transitory variance; temporary contracts

JEL classification: *J*21; *J*31

I. Introduction

A large and growing body 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 (which are related to changes in the quantity and prices of permanent individual characteristics) from a transitory component that captures the extent of earnings instability; see the recent review by Meghir and Pistaferri (2011).[1] The distinction between permanent and transitory inequality is important for various reasons. First, it is useful in evaluating the welfare implications of changes in inequality. An increase in permanent inequality would cer-

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[1] An incomplete list of the studies includes the following: Haider (2001) and Moffitt and Gottschalk (2012) for the US; Dickens (2000), Ramos (2003), and Alessie and Kalwij (2007) for the UK; Baker and Solon (2003) for Canada; Cappellari (2004) for Italy; and Bingley *et al.* (2013) for Denmark.

tainly reduce welfare, while increasing transitory inequality would have a weak effect on welfare measures, unless there are liquidity constraints that restrict consumption smoothing.[2] Second, it informs the assessment of different explanations for the increase in inequality: if rising inequality reflects an increase in permanent inequality, then a consistent explanation would be skill-biased technical change; in contrast, an increase in transitory inequality could reflect greater flexibility among workers in switching jobs, and therefore higher income mobility.

Tenure is a potentially important covariate of instability and, in this paper, we model their relation. Although we are not the first to model tenure in earnings variance models, we are the first to model the transitory component of earnings with respect to tenure.[3] Previous studies modeled the relation between permanent earnings and tenure, and were motivated by testing between alternative theories of wage determination (Parent, 2002). We believe that job tenure should also affect earnings instability. Specifically, we should expect earnings instability to decrease with job duration if there is employer-learning about the quality of the match over time (Lange, 2007), or if firms insure earnings against volatile shocks as they improve their knowledge of the match quality (Guiso *et al.*, 2005). The earnings model of this paper includes tenure effects on both the transitory and permanent components of the earnings process.

The recent body of literature has gone beyond the simple permanent–transitory earnings decomposition, using two different strategies. The first strategy involves reduced-form models, which examine the correlates of instability. Cameron and Tracy (1998) and Baker and Solon (2003) have explored the relation between instability and age. Although some previous studies have investigated the impact of both quits and layoffs on the transitory variance of wages in the US, none of these papers has explicitly modeled the effect of tenure on instability (see Huff Stevens, 2001; Hospido, 2012; Leonardi, 2012). A neighboring and growing body of literature focuses on earnings volatility (i.e., the variance of year-to-year

[2] Besides the distinction permanent–transitory, the distinction between "predictable" and "unpredictable" (Cuhna *et al.*, 2005; Cuhna and Heckman, 2007) and "insurable" and "uninsurable" (or partially insurable) shocks is particularly common in the macro literature (Heathcote *et al.*, 2010, 2014). This distinction requires the modeling of the structure of credit and insurance markets, and other risk-sharing characteristics of households (labor supply, family networks, etc.) or of government transfers, which might crowd out private transfers and self-insurance (Blundell *et al.*, 2015). For our purposes, it is enough to say that the variance of transitory shocks is very relevant for individual measures of welfare, either because it is unpredictable or because it is uninsurable.

[3] Many papers have focused instead on estimating the average returns to tenure: see Altonji and Williams (2005), Dustmann and Meghir (2005), and references therein.

earnings changes), and finds mixed evidence on the relation between workers' turnover and volatility (see Dahl *et al.*, 2011; Venn, 2011; Ziliak *et al.*, 2011; Celik *et al.*, 2012; Cappellari and Jenkins, 2014). However, papers that study earnings changes rather than levels do not distinguish between permanent and transitory shocks to earnings (Shin and Solon, 2011; Dynan *et al.*, 2012). A common problem with this literature on reduced-form models is, of course, the endogeneity of job-to-job mobility. We try to address both problems: we estimate the effect of tenure separately on the permanent and transitory components of earnings, and we identify changes in workers' tenure with cohort-time differences and with the differential exposure of birth cohorts to the succession of reforms of temporary employment contracts in Italy.

The second strategy, pursued in another strand of the literature, has analyzed the economic forces behind the degree of persistence and of variability in earnings, building structural models to better characterize behavior. Low *et al.* (2010) model labor supply and job mobility in a search and matching framework. Their approach is explicit about distinguishing between shocks and responses to shocks (i.e., job mobility) and between different types of uncertainty, loosely associated with employment risk (i.e., rates of arrival of job offers) and productivity risk (i.e., shocks to the match). Confirming that job-to-job mobility is important, they find that if mobility is ignored, the estimated variance of the permanent innovation to wages doubles, leading to an impression of much greater risk in the earnings process. This is because many of the wage fluctuations are due to individuals' moving to jobs with better match-specific effects. Flabbi and Leonardi (2010) and, on a more complex scale, Altonji *et al.* (2013) estimate a model of wages and transitions between jobs and into unemployment driven by exogenous shocks, which are the underlying source of fluctuations. In the same line of research but with more focus on wage growth, Adda *et al.* (2013) model workers' career progressions in a framework in which wages grow because workers learn on the job and through job shopping. Relative to reduced-form estimates, these models have the advantage that one can construct counterfactual life-cycle profiles, by comparing profiles with and without returns to experience, tenure, or job mobility. However, this is obtained at the cost of imposing specific distributional assumptions.

We use Italian Social Security records between 1986 and 2003, and illustrate how the variation in tenure is connected with a succession of labor market reforms that liberalized temporary employment contracts and had a differential impact across birth cohorts and over time. In the US, it has been difficult to establish a link between earnings instability and workers' tenure because the empirical literature has found little evidence of

a decline in average tenure data.[4] Unlike the US, in Italy and in many other continental European countries, the diffusion of temporary employment contracts (i.e., short-tenure contracts) has generated additional variation in tenure, and constitutes an "institutional" reason for shorter tenure. We exploit the differential incidence of temporary contracts across cohorts and over time to characterize the relation between earnings instability and tenure.

Our approach is not structural but we contribute to the earnings dynamics literature by proposing a tractable method to estimate the effect of tenure on instability, based on the use of between-cohort variation. We begin with a model in which earnings instability depends over time on cohort-specific average tenure. We control for time and cohort fixed effects in both the transitory and the permanent components of the earnings process, so that the effect of tenure on instability is estimated in a difference-in-differences set-up. Next, we use the same set-up to estimate the effect of temporary employment on earnings instability, a kind of reduced-form estimate of the impact of reforms. In our final analytical step, we project tenure on cohort-specific temporary employment, and use this prediction in place of actual tenure within the earnings dynamics model with cohort and time fixed effects. In this way, we explicitly model the idea that tenure is a function of policies that has affected the diffusion of short-term contracts differentially across birth cohorts.

We find that tenure reduces earnings instability by between 11 and 13 percent per year, depending on whether actual or predicted tenure is used in the earnings dynamics model. The decline of instability is not constant but it is faster in the initial years of the match: this finding is consistent with employer-learning effects (e.g., Lange, 2007). When we look directly at the effect of temporary contracts, we find that instability is the largest in cohorts with a high incidence of temporary employment, and the gap with other cohorts has been widening after the most recent waves of labor market reform that have liberalized temporary employment in the late 1990s. We can exclude the fact that this result stems from the selection of more stable workers into more stable jobs because our model exploits variation of tenure and instability across birth cohorts and over time, while controlling for heterogeneity with cohort and year fixed effects. Any remaining selection occurring within cohorts (such as the individual choice of being a temporary employee) is irrelevant because it is not the source of variation used in the estimation.

[4] Jaeger and Huff Stevens (1999), Gottschalk and Moffitt (1999), and other contributions in the same special issue of the *Journal of Labor Economics* find little evidence of a decrease in workers' tenure. More recently, Farber (2010) finds some decrease in the average tenure of older workers.

Finally, our paper is also of interest to the large body of literature that has studied the effect of temporary contracts on employment, job flows, and wage levels, because nobody has looked so far at their effects on earnings instability.[5] Yet, one of the main policy concerns about the diffusion of temporary employment contracts is their implications in terms of earnings instability and welfare, because the temporary part of earnings variance is often uninsurable in the presence of imperfect capital markets and liquidity constraints. In this paper, we fill in the gap and provide an estimate of the earnings instability directly associated with a temporary contract: young workers in cohorts with a high incidence of temporary contracts have an earnings instability between 50 and 100 percent higher than workers who belong to cohorts with a low incidence of temporary contracts. We believe that these results might be useful from the policy point of view, because they highlight new channels through which temporary employment can affect individual well-being.

The rest of this paper proceeds as follows. In Section II, we describe the data. In Section III, we describe the institutional background and the concurrent evolution of tenure, temporary employment, and earnings instability. In Section IV, we lay out the error component models of the impact of tenure and temporary contracts. In Section V, we present the results, and we conclude in Section VI.

II. Data

The data are drawn from the Italian Social Security Administration (*Istituto Nazionale di Previdenza Sociale*, INPS) archives, and span the years 1985–2003. 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 dataset contains individual longitudinal records generated using social security numbers. However, because 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, or retirement. This selection is common for administrative data, which typically include the private sector only. To provide some information on Italian private sector employment in comparison to other sectors of the economy, we can use external data sources. Us-

[5] Temporary jobs are known to pay less, offer less training, and be less satisfying than regular jobs (Booth *et al.*, 2002). The evidence on whether temporary jobs are stepping stones to permanent jobs is mixed (Autor and Houseman, 2012).

ing the Bank of Italy data (Survey of Households Income and Wealth, SHIW) for 1998, it appears that the private sector constitutes 52 percent of total employment, agriculture represents only 2 percent, while public employment and self-employment represent 23 percent each.[6] 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 transitions into other states show that workers are very stable in the private sector. After two years (always using SHIW data), 83 percent of male workers aged between 21 and 55 employed in the private sector in 1998 are still working in the private sector, 7.5 percent moved to the public sector, only 3 percent to self-employment, 2.3 percent to unemployment, and 2.5 percent to retirement.

As is common with administrative data sources, the amount of observable individual characteristics in the INPS data is limited. We have information on employees' age, gender, occupation (blue collar/white collar), yearly earnings, number of paid weeks, the initial and final months of job matches, and the type of contract (permanent/temporary).

Sample Selection

The administrative data in electronic form start in January 1985, and the start date of all contracts already running at that date is artificially set to January 1985. In order to measure tenure accurately, we consider only matches starting after January 1, 1985. Because such a selection rule leaves only a few observations in 1985 compared to the other years in the panel, we consider data from 1986 onwards. We keep in the sample all male workers aged 21 to 55 with positive earnings, who work as blue collar or white collar workers in the non-agricultural private sector between 1986 and 2003. The selection on age is aimed at avoiding the extremes of the working career, because employment volatility just after entry into the labor market or close to retirement might blur the measurement of earnings instability.[7] As customary in this literature, we focus on males because their labor force participation is less endogenously intermittent than that of females.

Previous studies, such as Haider (2001), Baker and Solon (2003), and Moffitt and Gottschalk (2012), have demonstrated the existence of relevant age and calendar time effects in both the permanent and the transitory components of earnings. Because we estimate tenure effects over a long

[6] While there is evidence that wages are less volatile in the public sector than in the private sector (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.

[7] As a robustness check on the age selection, we alternatively used 25 and 30 as starting points for the age–earnings profile, and found the results of the analysis to be unaffected.

period, it is crucial to control for age and time effects. To disentangle the two effects within our econometric model of earnings dynamics, we form subsamples defined by the year of birth (birth cohorts), and use them jointly in estimation. In order to ease the identification of age–earnings profiles within each cohort, we set the minimum length of observation of a cohort to ten years. Given our sample selection on age, this implies that we consider cohorts of individuals born between 1940 (who turn 55 in 1995, in the tenth year of data in the sample) and 1973 (who turn 21 in 1994, and can be observed ten times before the end of the sample). Cohorts born between 1948 and 1965 are observed 18 times (i.e., over the whole sample period), while for cohorts born before 1948 or after 1965, the number of data points monotonically decreases, going from 17 for those born in 1947 and 1966 to ten for the oldest and youngest cohort born in 1940 and 1973. There are 34 birth cohorts in total. It needs to be stressed that besides allowing the identification of time and cohort effects, the cohort-by-year variation is important in our paper as it provides us with variation in tenure and incidence of temporary contracts, which are two key variables in our models (see the next section).

In the course of the paper, we use real weekly earnings (yearly earnings in 2002 prices divided by the number of weeks paid). For the cases of multiple job spells in the same year, we consider the longest spell. In order to reduce the influence of outliers, we drop the top and bottom three observations from the cohort-specific yearly wage distribution. We also exclude individual earnings histories characterized by excessive churning (which might inflate the measurement of earnings instability), and require for each individual a minimum of five consecutive earnings observations, a selection rule that is intermediate between the one used by Baker and Solon (2003), that is, continuous earnings strings, and the approach of Haider (2001), who allows individuals to move in and out of the sample with the only requirement of having two valid but not necessarily consecutive observations on earnings.

Descriptive Statistics

The dataset resulting from this selection includes 48,226 individuals with job spells that started after January 1, 1985, and at least five consecutive years of valid wages yielding 552,209 person–year observations over the years 1986–2003. The truncation of the sample (and the consequent dropping of matches that started before January 1, 1985) is a common procedure with administrative data that do not contain information on tenure at entry. The corresponding sample without truncation would include approximately 58,000 individuals and 720,000 person–year data points, which raises the issue of whether the truncation of the sample induces some se-

Table 1. *Descriptive statistics*

Year	No. obs.	Age	Mean of log wages	Std dev. of log wages	Percent temporary	Tenure
1986	13,358	32.33	5.84	0.37	0.05	10.24
1987	16,523	32.46	5.86	0.37	0.08	17.45
1988	19,641	32.78	5.87	0.37	0.09	23.56
1989	22,227	33.22	5.91	0.37	0.10	28.85
1990	24,496	33.66	5.94	0.38	0.09	33.35
1991	27,970	34.10	5.96	0.39	0.06	36.92
1992	29,328	34.66	5.97	0.39	0.05	42.47
1993	35,432	35.64	5.97	0.40	0.04	38.71
1994	36,443	36.23	5.95	0.40	0.04	44.70
1995	37,164	36.84	5.96	0.41	0.04	48.23
1996	37,355	37.37	5.97	0.41	0.04	52.61
1997	37,650	37.93	6.00	0.42	0.04	57.06
1998	37,340	38.52	6.01	0.42	0.07	61.07
1999	38,002	38.93	6.01	0.44	0.07	62.27
2000	37,044	39.66	6.01	0.44	0.06	66.84
2001	35,723	40.37	6.03	0.44	0.05	71.57
2002	34,133	41.01	6.04	0.45	0.05	75.73
2003	32,380	41.62	6.04	0.45	0.05	79.84

Notes: $N = 48{,}226$; $NT = 552{,}209$. The sample contains only matches starting after January 1, 1985.

lection bias. We address these concerns in our background working paper (Cappellari and Leonardi, 2013), where we provide estimates of earnings models on both the truncated and non-truncated samples, showing that sample selection bias is not an issue.

Table 1 shows descriptive statistics of the sample. The yearly number of observations increases in the initial years because, as discussed above, we consider only new matches. Average age increases over time at a rate of approximately six months per year, due to the revolving-by-cohort design of the sample, which separates time and age trends. The table also shows real wage growth (approximately 1 percent on a yearly basis) and increasing dispersion.

We use all valid wage observations in our sample to estimate the covariance structure of earnings for the 34 birth cohorts. While not solving issues of endogenous panel attrition, such an unbalanced panel design is certainly less restrictive than analyses based on balanced panels.

We plot estimated variances and covariances averaged across birth cohorts in Figure 1. The earnings dispersion appears to increase at a steady pace over the period. These patterns reproduce the evidence for Italy provided by other studies (e.g., Brandolini *et al.*, 2002). The covariances at various lags are at a lower level than the variance, but still show an upward trend. As expected, the distance between the covariances at increasing lags decreases over lags, and the covariances tend to stabilize to a long-term

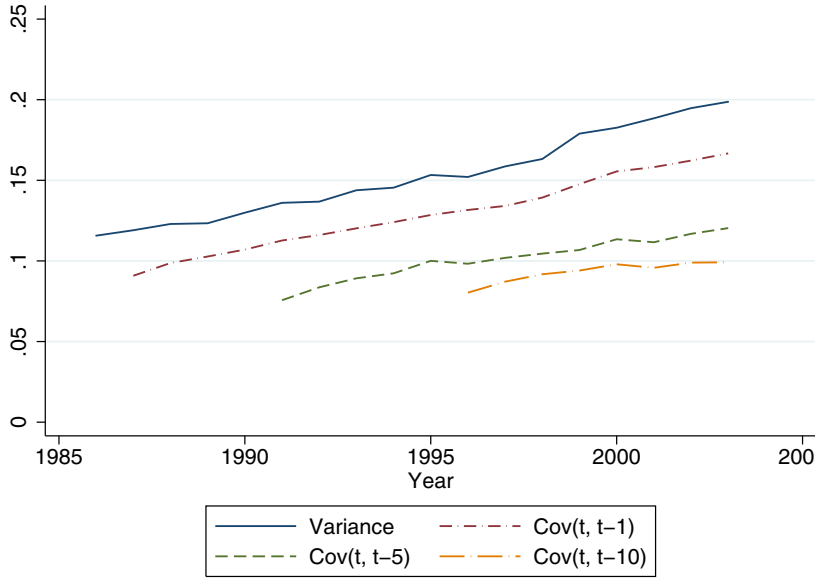

Fig. 1. Earnings variances and covariances

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 a low-order autoregression (Gottschalk and Moffitt, 2009). Both elements feature in our model of Section IV.

Table 1 also shows the incidence of temporary contracts by year. We now discuss the link between these trends and the reforms in the labor market.

III. Institutional Background and Patterns of Instability 1985–2003

As with other European countries, labor market flexibility has increased in Italy over the last 20 years, through a series of measures that have introduced various types of temporary contracts without changing the legislation on permanent, open-ended contracts, thus affecting mostly new entrants into the labor market.

We exploit, in particular, two waves of reforms occurring during the sample period. The first wave of reforms of temporary employment contracts took place in the mid-1980s with the introduction of "work and training" contracts (*contratti di formazione e lavoro*). These are fixed-term employment contracts used in particular to hire young non-manual workers.

They became very popular for various reasons. First, using this contract, employers pay lower social security contributions and, in exchange, provide some training. Secondly, firms pay no dismissal costs when these contracts expire. Finally, the main advantage of a work-and-training contract relative to an apprenticeship contract is that all training is supposed to take place within the firm rather than with external bodies (which has always been considered a nuisance by firms). For all these reasons, the work-and-training contracts saw a large diffusion in the late 1980s. However, their misuse by employers – in particular, the failure to provide the required training – gave rise to much litigation, which in turn led to a reduced use of these contracts through the mid-1990s, and finally to their abolition in the early 2000s.

The second wave of reforms took place in the late 1990s. In 1997, the Treu Package (named after the, then, Minister of Labor) legalized temporary work agencies and liberalized fixed-term contracts (which are both forms of temporary employment contracts). Since that time, agency workers are typically more expensive than workers hired with a standard open-ended contract but they can be dismissed at will, while fixed-term employment contracts have a legal duration of two years, and can be renewed only once.

Table 1 reports the average incidence of temporary employment contracts (the sum of employees with a work-and-training contract, fixed-term employees, and agency workers). Temporary contracts are more widespread when there is a reform (i.e., in 1988, when there was the peak use of work-and-training contracts, and ten years later with the introduction of the Treu reform). However, this average masks a great deal of variation in the percentage of temporary employment contracts by cohort, because they were mainly used to hire young labor market entrants. In the estimation, we exploit this variation and its relation to average tenure.

Temporary Employment and Tenure in the Years of the Labor Market Reforms

Table 2 shows the average tenure, in months, for workers in permanent and temporary contracts within selected cohorts, as well as the incidence of temporary employment over time. We select young cohorts because they have a higher incidence of temporary contracts. All cohorts observed since the beginning of the panel start with low average tenure because the average refers only to contracts that started after January 1985. Considering the cohorts born in 1965, 1970, and 1973 at the tenth year of observation (1995, 2000, and 2003, respectively), we can see that the accumulation of tenure on permanent contacts is similar across cohorts: older cohorts accumulate, on average, a slightly longer tenure as a result of the lower job mobility of older workers relative to younger ones.

Table 2. *Average tenure in months and temporary employment*

Year	Tenure of permanent employees	Tenure of temporary employees	Share of temporary employees
Cohort born 1965			
1986	9.34	6.67	0.17
1987	16.23	10.80	0.25
1988	22.68	12.87	0.26
1989	28.48	15.21	0.22
1990	33.22	14.16	0.17
1991	36.44	14.89	0.13
1992	41.69	16.39	0.09
1993	41.53	14.52	0.06
1994	47.88	14.77	0.06
1995	52.70	17.98	0.05
1996	56.66	19.22	0.06
1997	60.78	18.59	0.05
1998	66.54	23.87	0.06
1999	68.60	20.10	0.06
2000	72.16	20.18	0.06
2001	76.11	25.33	0.06
2002	80.04	21.42	0.06
2003	85.49	21.83	0.07
Cohort born 1970			
1991	20.04	11.94	0.25
1992	23.23	14.96	0.21
1993	26.12	15.64	0.18
1994	31.87	14.81	0.14
1995	35.44	16.49	0.12
1996	39.40	14.87	0.13
1997	43.43	15.55	0.13
1998	48.00	18.09	0.18
1999	48.92	20.11	0.17
2000	54.03	26.58	0.16
2001	58.25	29.47	0.13
2002	63.64	22.99	0.12
2003	68.71	29.17	0.11
Cohort born 1973			
1994	24.32	9.34	0.16
1995	25.76	11.97	0.21
1996	28.00	14.05	0.20
1997	31.89	15.83	0.16
1998	37.37	16.36	0.24
1999	39.46	15.92	0.23
2000	44.37	19.46	0.22
2001	49.61	23.09	0.20
2002	55.10	22.79	0.19
2003	59.66	27.42	0.18

Table 2 also shows the incidence of temporary employment contracts and the average tenure of temporary workers for the same cohorts. The overall share of temporary contracts reached 26 percent in 1988 for the cohort born in 1965, in the peak year of the diffusion of work-and-training

contracts. In 1997, the Treu reform liberalized temporary employment, and our data show that its incidence increased substantially between 1997 and 1998 for the youngest cohorts (18 and 24 percent is the incidence of temporary contracts in the year 1998 for the cohorts born in 1970 and 1973, respectively). Table 2 clearly shows that while permanent workers accumulate tenure on the job, the average tenure of temporary workers is always below 30 months. This table shows that the incidence of temporary contracts varies from cohort to cohort, and is also related to the average tenure of a given cohort in a given year, illustrating the sources of variation that we use to estimate the earnings dynamics model that is laid out in the next section.

Patterns of Earnings Instability and Tenure

An initial description of the relation between earnings instability and tenure in the reform years can be provided using the method proposed by Gottschalk and Moffitt (1994), and by relating individual-level transitory fluctuations in earnings with individual tenure or type of employment contract. We take data within a time window $[t-q, t+q]$ and consider individuals with continuous earnings strings within that window. The mean of each individual's earnings within the time window constitutes his permanent component of earnings, while deviations from the mean represent the transitory earnings component. The individual-level variance of deviations from the mean is our measure of earnings instability,

$$\sigma_{it}^2 = \frac{1}{2q} \sum_{s=t-q}^{t+q} (w_{is} - w_i)^2, \tag{1}$$

where w_i is the average of the individual's log wages within the window $[t-q, t+q]$. Repeating this calculation while moving the fixed-length window forward provides a series of measures of earnings instability for each individual.[8]

We describe the relation between instability and tenure by regressing the instability measure on workplace tenure, dummies for year, occupation, age, region, firm size, and one digit industry, plus individual fixed effects. Workplace tenure enters these descriptive regressions either linearly (measured months) or through linear splines (with knots at one, two, three, and four years). Alternatively, we assess the effect of employment stability

[8] We present the results for $q = 2$ (i.e., using time windows of five years); we find the results to be robust when using $q = 4$.

Table 3. *Impact of tenure and temporary contracts on earnings instability: Gottschalk and Moffitt (1994) method*

	Specification 1		Specification 2		Specification 3	
	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Months of tenure	−0.0002	0.0000				
0< tenure <= 12 months			−0.0040	0.0002		
12< tenure <= 24 months			−0.0024	0.0001		
24< tenure <= 36 months			−0.0013	0.0001		
36< tenure <= 48 months			−0.0004	0.0000		
48< tenure			−0.0001	0.0000		
Time in temp. contracts					0.0499	0.0014
Constant	0.0743	0.0226	0.1359	0.0225	0.0680	0.0226
R^2	0.0293		0.0727		0.0166	

Notes: $N = 48{,}127$; $NT = 315{,}484$. The results for specification 1 are estimated from the equation, $\sigma_{it}^2 = X_{it}\beta + \delta ten_{it} + u_i + e_{it}$, where individual controls X_{it} include a set of year, occupation, age, region, firm size, and one digit industry dummies, and u_i is an individual fixed effect. In specification 2, ten_{it} is substituted with a spline. In specification 3, we instead use time in temporary contracts, which refer to the percentage of years that each individual has spent in temporary contracts within a five-year window.

by looking at temporary contracts, specifically the share of time spent on temporary contracts within the five-year window.

The results of this exercise are provided in Table 3 and show a clear decline of instability with tenure. The first specification shows a significant decline of instability with each month of tenure; the second specification with tenure splines confirms this result. The third specification shows that instability increases with the average time spent on temporary contracts.

The approach of Gottschalk and Moffitt (1994) is very simple and intuitive, and generates individual-specific measures of instability that, instead, are not available when using more formal models. However, it has disadvantages. First, it assumes constant permanent earnings and white noise transitory earnings, while the earnings dynamics literature has shown both assumptions to be unrealistic because of life-cycle effects and serially correlated shocks. Secondly, the method does not necessarily pick up the exact turning points in the time series of transitory variances when the turning points fall within the time windows used for the averaging of individual earnings. We overcome both types of limitations with the formal model of the next section.

IV. Models of Earnings Instability and Tenure

In this section, we develop an econometric model of earnings dynamics and instability that takes into account the effects of workplace tenure while

controlling for both serial correlation in transitory earnings shocks and life-cycle effects in permanent earnings. Both are important features of the earnings process, which were ignored in the simplistic approach that we used at the end of the previous section. In Section III we documented a large variation in the incidence of temporary employment by birth cohorts and a concurrent variation in tenure, arguing that these patterns were induced by the labor market reforms taking place over the period. Now, we exploit this variation and model earnings instability as a function of cohort-specific average tenure over time. In this way, our estimates will not reflect unobserved heterogeneity within birth cohorts. We will ensure that the estimated effects do not reflect any omitted heterogeneity between cohorts by controlling for time and cohort effects throughout the model, so that selection effects operating between cohorts over time will be controlled for in a difference-in-differences set-up.

First, we introduce tenure effects into a model of life-cycle earnings where instability is defined as the population variance of transitory earnings shocks. Next, we alternatively characterize individual earnings dynamics in terms of the type of employment contract, temporary versus open-ended, a type of reduced-form estimation of the effects of reforms. Finally, we use the variation in temporary contracts across cohorts and time periods to predict cohort-level job tenure over time, and we use this predicted measure in place of the actual one within the model of earnings instability and tenure. This latter strategy makes explicit the link between variation in temporary employment (in turn driven by labor market reforms) and variation in tenure.

We specify our models in terms of the log-earnings deviations from period- and cohort-specific means. The removal of the period- and cohort-specific means is equivalent to controlling for cohort-specific age fixed effects, which is crucial in our context as we are interested in individual life-cycle profiles, which might be confounded by cohort-specific wage growth over the life cycle (Baker and Solon, 2003). Other papers in the literature have used first-stage regressions that include on the right-hand side polynomials in age, cohort effects, and other individual characteristics (e.g., Meghir and Pistaferri, 2004; Moffitt and Gottschalk, 2012). Empirically, we obtain de-meaned log-earnings as residuals from cohort-specific regressions on time dummies. Individual log-earnings deviations w from the period- and cohort-specific means are the sum of a permanent (long-term) component y and an orthogonal transitory shock v and orthogonality holds by definition of the permanent and transitory components of earnings:

$$w_{it} = y_{it} + v_{it}; \qquad E(y_{it}v_{it}) = 0. \tag{2}$$

Here, $i = 1, \ldots, N$ denotes the individuals, $t = t_c, \ldots, T_c$ denotes the time periods, and $c = c(i)$ denotes the birth cohort of person i. The observation window is cohort-specific, due to the revolving-by-cohort design of the sample.

Modeling the Impact of Tenure

We extend existing specifications (e.g., by Baker and Solon, 2003) by including tenure effects. Models in the literature typically allow for life-cycle variation in permanent earnings plus mildly persistent mean reverting shocks (the transitory component). We characterize the evolution of earnings dispersion occurring within job matches on top of life-cycle variation between job matches. Match-specific effects affect both the permanent and the transitory components, and each becomes a function of workplace tenure. Shocks to the permanent component have a unit root and capture any permanent reshuffling occurring to the earnings distribution within job matches (e.g., because of promotions). Effects on the transitory component (our proxy of earnings instability) are captured through the variation in transitory shocks dispersion with tenure. If such dispersion is a proxy for measurement error on the underlying productivity of the match, we should expect its size to decline with tenure as information on match quality is revealed over time.

Our specification of permanent earnings is as follows:

$$y_{it} = (\alpha_i + \beta_i Age_{it} + S_{it} r_{it})\pi_t \lambda_c; \qquad r_{it} = r_{it-1} + q_{it}; \tag{3}$$

$$q_{it} \sim (0, \sigma_q^2); \qquad (\alpha_i, \beta_i) \sim (0, 0; \sigma_\alpha^2, \sigma_\beta^2, \sigma_{\alpha\beta}). \tag{4}$$

Permanent earnings vary both over the life cycle and within job matches. Life-cycle variation is parametrized via the so-called random growth (RG) model, which assumes individual-specific linear earnings profiles in age or experience (we use age). This simple specification can capture important features of individual earnings dynamics. The model allows for the heterogeneity of permanent earnings, both at the beginning of the working life (through the variance of the intercepts σ_α^2) and over time (through the variance of the slopes σ_β^2). Models such as those of Mincer (1958) and Ben-Porath (1967) predict that human capital investments will induce a trade-off between starting earnings levels and earnings growth, as investors give up some of their earnings at the beginning of their working career and increase their productivity over the career. Therefore, these models predict that the covariance between the intercepts and slopes of individual-specific profiles ($\sigma_{\alpha\beta}$) will be negative. However, the two parameters α_i and β_i

might simply reflect different dimensions of individual ability – say, the ability to accumulate human capital both prior to labor market entry and on the job – in which case one would expect a positive covariance between the intercepts and slopes of the profile (Gladden and Taber, 2009).[9]

We augment the life-cycle dynamics with a unit root job-specific shock r_{it} – a random walk (RW) process – which affects the earnings dynamics of job stayers, with S_{it} being a dummy for stayers between periods $t-1$ and t. This captures persistent earnings differences within job matches. Therefore, the overall specification of the permanent earnings component is an RG plus an RW. The model parameters are identified by quadratic age trends (RG) and linear tenure trends (RW) in the intertemporal covariance structure of earnings. We separate life-cycle effects from time trends by exploiting cohort-specific earnings covariance structures, which generates a variation across cohorts and time periods in the intertemporal earnings distribution. Therefore, RG–RW parameters are identified by variation in tenure and age across birth cohorts over time. Period and cohort loadings, π_t and λ_c, take into account the aggregate shifts in the long-term distribution of earnings operating between cohorts over time, ensuring that the estimated RG–RW parameters will not reflect cohort or time trends.[10]

We model transitory earnings as an AR(1) process with cohort-specific effects, period-specific effects, and a variance of shocks that depends on both age and tenure.[11] The latter allows the estimation of the effects of tenure on earnings instability. Baker and Solon (2003) specified the variance of transitory shocks as a quartic in age, and exploited the variation in age across cohorts and time periods for estimation, finding significant age variation. Here, we use more flexible exponential splines. Our exponential specification ensures non-negativity, while preserving flexibility through the spline function. More importantly, we extend their approach to include

[9] See Cappellari (2004) for evidence of positive intercept–slope covariance in the INPS data.
[10] In the macro labor literature, the RG model is also called the heterogenous income profile (HIP), while the RW model is known as the restricted income profile (RIP), and there is a debate on their performance in fitting idiosyncratic trajectories of labor incomes. Guvenen (2007) and Hryshko (2012), among others, discuss the two parametrizations in the context of models of life-cycle optimization with rational expectations. In this paper, we use a mixture of the two processes to capture earnings dynamics within job matches and over the life cycle. The RG plus RW specification is also used in Baker and Solon (2003) and Moffitt and Gottschalk (2012), but while in their papers both processes evolve only over age, in our case the two processes evolve along two different dimensions: age and workplace tenure.
[11] We have also experimented with ARMA(1,1) specifications. However, when we model the impact of tenure on instability, moving average components are difficult to identify. 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.

tenure effects in the variance of shocks. Our transitory earnings model is written as

$$v_{it} = \tau_t u_{it} = \tau_t(\rho u_{it-1} + \varepsilon_{it}); \qquad \varepsilon_{it} \sim (0; \sigma^2_{\varepsilon ct}); \qquad u_{i0} \sim (0; \eta_c \sigma^2_0). \quad (5)$$

Here, τ_t are period-specific factor loadings and u_{i0} is the initial condition for person i, whose variance is cohort specific. Period- and cohort-specific shifters τ_t and η_c control for aggregate shifts in the distribution of transitory earnings. Our measure of earnings instability is the variance of AR(1) innovations and depends on both age and tenure (Ten_{ct}), averaged across cohorts and time periods,

$$\sigma^2_{\varepsilon ct} = \sigma^2_{\varepsilon} \exp[g_1(Age_{ct}) + g_2(Ten_{ct})], \quad (6)$$

where the spline functions g_1 and g_2 parametrize the evolution of earnings instability over the life-cycle and within job matches. Equation (6) connects earnings instability with tenure, and represents the main contribution of our model. It is important to stress that, besides controlling for time and cohort fixed effects, we estimate tenure effects holding constant life-cycle trends. This rules out the possibility of differential selection operating between cohorts over time due to age differences when the cohorts are exposed to changes in labor market institutions.

Modeling the Impact of Temporary Contracts

As discussed in Section III, much of the variation in tenure comes from the diffusion of temporary contracts. An alternative way to measure the relevance of tenure for earnings instability is to look at the type of contract: open-ended or temporary. The underlying idea is that temporary contracts are associated with job turnover and do not favor the accumulation of seniority. Thus, if tenure reduces instability, then we should expect larger instability on temporary contracts relative to open-ended contracts.[12]

As was the case with tenure, we model the impact of contract types on instability by exploiting the variation in the incidence of temporary contracts across cohorts and time periods, allowing for time and cohort effects, to capture unobserved heterogeneity along those dimensions. Any remaining selectivity operating within-cells is irrelevant to our results because that variation is not used in the estimation.[13]

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[12] Clearly, such a "reduced form" could pick up various other reasons why temporary contracts affect instability. For example, temporary contracts might lower the commitment of employers and employees, and reduce training (Booth *et al.*, 2002).

[13] The issue of within-cohort selection into temporary contracts might be important. If, for example, the least stable workers are the first to be offered temporary contracts, then temporary contracts just act as a mechanism to sort workers into those who generally have short tenure and those who tend to stay longer in the same job. While, overall, this would

We define two dummies for the incidence of temporary contracts in total employment being between 5 and 10 percent, or above 10 percent: $F1_{ct}$ and $F2_{ct}$. We specify the variance of shocks as a function of the two dummies and the age spline, so that equation (6) is replaced by the following expression:

$$\sigma^2_{\varepsilon ct} = \sigma^2_{\varepsilon} \exp[g_1(Age_{ct}) + \phi_1 F1_{ct} + \phi_2 F2_{ct}]. \quad (7)$$

For long-term earnings, we take an approach similar to the one used for instability, and allow their variance to be a function of the incidence of temporary contracts over cohort–period cells. The type of employment contract should also affect the permanent part of the earnings process: insofar as temporary contracts are less favorable to the development of job-specific skills and are characterized by less training than open-ended contracts, we expect the distribution of long-term earnings to be more compressed among temporary workers. Therefore, the total variance of permanent earnings will depend on the incidence of temporary contracts, a quadratic in age (through the RG specification) and cohort and period factor loadings:

$$\text{var}(y_{it}) = (\sigma^2_{\alpha} + 2Age_{ct}\sigma_{\alpha\beta} + Age^2_{ct}\sigma^2_{\beta})\pi^2_t\lambda^2_c \exp(\gamma_1 F1_{ct} + \gamma_2 F2_{ct}). \quad (8)$$

The parameters ϕ and γ measure the effects of temporary contracts on the transitory and the permanent components.

Modeling the Impact of Tenure through Variations in Temporary Contracts

We have shown in Section III that some of the variation in tenure occurs between individuals with different contractual arrangements, and that variation in the incidence of temporary employment contracts was more pronounced for young cohorts as a result of institutional changes over the period. In this respect, equation (7) can be seen as a reduced-form model in which instability is conditioned on the institutional source of variation in tenure. We now further exploit this idea and explicitly use the variation of temporary contracts between cohorts over time to estimate the effect of workplace tenure on earnings instability. In particular, first we use the cohort-level variation in temporary contracts to predict workplace tenure by cohorts over time. Next, we use this predicted measure in place of actual tenure in the earnings model. In this way, we exploit only the tenure

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have no impact on earnings instability, it would increase the instability for those on temporary contracts, and decrease it for those on permanent contracts. In this respect, the results of the simple instability model of Table 3 are reassuring because they show that the relation between instability and temporary employment is not driven by individual fixed effects.

variation that comes from variation in temporary contracts, and not from other time-varying unobservables between cohorts. As long as differential exposure to temporary contracts between cohorts can be considered exogenous, this strategy helps cope with spurious associations between tenure and earnings instability operating between cohorts that are not captured by the cohort and time shifters and that might undermine the estimation of the model. Our first stage regression is

$$Ten_{it} = \delta_0 + \delta_c d_{ic} + \delta_t d_{it} + \delta_F F_{ct} + e_{it}, \tag{9}$$

where Ten_{it} is the level of tenure for person i in year t, d are cohort and time fixed effects, F_{ct} is the proportion of individuals on temporary contracts in cohort c and period t, and e_{it} is a white noise error term.[14] Let $\widehat{Ten_{it}}$ denote the prediction from this model, which varies only at the time and cohort level. Such a variation occurs both because of cohort and time fixed effects (that are already controlled for in the earnings model) and because of changes in temporary contracts by cohort over time, which is the source of variation that we exploit for estimating the effect of tenure on earnings. We do this by replacing actual tenure with predicted tenure in the equation for earnings instability (equation (6)). Similarly, we replace actual by predicted tenure in the RW component of the model for permanent earnings. In each case, we take into account the fact that $\widehat{Ten_{it}}$ is an estimated regressor by using weights that are proportional to the inverse of its estimated variance.

V. Results

To set the scene, we begin our discussion by considering Figure 2, which plots the earnings variance decomposition into long-term and transitory components obtained using a model without tenure or contract type effects.[15] The predicted total variance of earnings replicates quite closely the patterns of the raw variance displayed in Figure 1, indicating that the fitting performance is good. These patterns suggest that the increasing earnings inequality in the late 1980s and early 1990s was 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

[14] We allow for unrestricted autocorrelation of the error over time by cohort using clustered standard errors at the cohort-by-year level. The results are unaffected if we use a specification with dummy variables, $F1_{ct}$ and $F2_{ct}$, in place of the cohort-specific proportion F_{ct}.

[15] We estimate the model parameters using the equally weighted minimum distance estimator (EWMD), which matches the moment restrictions generated by the earnings model with empirical moments obtained from the data. We use a robust variance estimator $\text{Var}(\theta) = (G'G)^{-1}G'VG(G'G)^{-1}$, where θ is the parameter vector, V is the matrix of fourth moments, and G is the gradient matrix evaluated at the solution of the minimization problem.

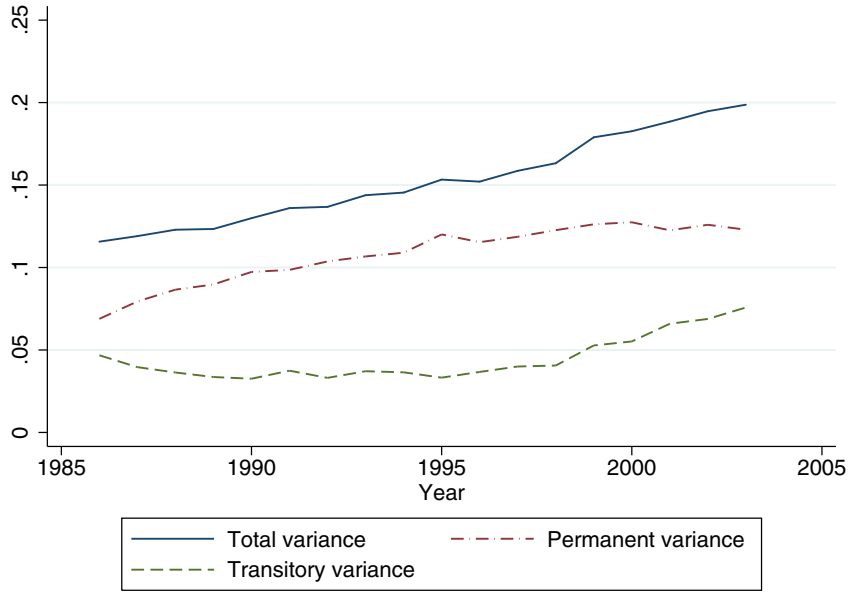

Fig. 2. Predicted variance components

technical change. The trends in the last part of the period have a different nature: while the growth of permanent inequality levels off after 1995–1996, earnings instability displays an upward pattern over the last years of observation, consistent with the increased labor market flexibility brought about by labor market reforms in this period.

Results of the Model with Tenure

Parameter estimates for the model, which includes tenure, laid out in equations (3)–(6), are presented in Column 1 of Table 4. The table provides estimates of "core" earnings components, while the estimated time and cohort shifters are presented in the Appendix. The parameter estimates for the permanent component show that the RG coefficients are precisely estimated, and indicate substantial heterogeneity of both initial earnings σ_α^2 and life-cycle earnings growth σ_β^2: an individual whose growth parameter β_i is located one standard deviation above the mean of the distribution of the growth rates experiences an earnings growth that is 1.4 percent ($\sqrt{0.0002} \times 100$) faster than the mean. Moreover, the covariance between the intercepts and the slopes of the RG is positive ($\sigma_{\alpha\beta} > 0$), which indicates that the sources of earnings heterogeneity at the start of the working career and over the career complement each other, which might reflect heterogeneous abilities in human capital accumulation affecting both schooling

Table 4. *Models of earnings dynamics: estimates of RG, RW, and AR(1) parameters*

	(1) Model with tenure		(2) Model of temp. contracts		(3) Model with predicted tenure	
	Coeff.	SE	Coeff.	SE	Coeff.	SE
Permanent component						
σ_α^2	0.0167	0.0010	0.0211	0.0014	0.0164	0.0010
$\sigma_{\alpha\beta}$	0.0021	0.0001	0.0021	0.0001	0.0021	0.0001
σ_β^2	0.0002	0.00002	0.0002	0.00003	0.0002	0.00002
σ_q^2	0.0015	0.0010			0.0022	0.0009
γ_1			−0.0996	0.0181		
γ_2			−0.0950	0.0438		
Transitory component						
σ_ε^2	0.0489	0.0124	0.0600	0.0093	0.1286	0.0782
σ_0^2	0.0575	0.0116	0.0738	0.0122	0.0567	0.0115
ρ	0.4236	0.0200	0.5611	0.0086	0.3875	0.0211
g_{11}	0.0468	0.0374	−0.0002	0.0151	0.0885	0.0356
g_{12}	0.0283	0.0176	0.0064	0.0125	0.0389	0.0191
g_{13}	0.0362	0.0141	0.0033	0.0098	0.0492	0.0150
g_{14}	−0.0968	0.0087	−0.0975	0.0074	−0.0938	0.0086
g_{15}	0.1309	0.0140	0.0959	0.0110	0.1284	0.0131
g_{21}	−0.3892	0.2968			−1.1581	0.4683
g_{22}	−0.3167	0.1511			−0.5050	0.1567
g_{23}	−0.3548	0.0963			−0.4801	0.1090
g_{24}	−0.1616	0.0733			−0.2212	0.0854
g_{25}	−0.1950	0.0512			−0.2422	0.0535
ϕ_1			0.1417	0.0475		
ϕ_2			0.2307	0.0723		

Notes: $N = 48{,}226$; $NT = 552{,}209$. The models include time and cohort shifters in both the permanent and transitory components whose estimates are reported in the Appendix. The models are estimated on 4,686 earnings moments over the period 1986–2003, and on 34 birth cohorts born between 1940 and 1973. The g_1 coefficients refer to the spline in age with knots at 26, 31, 36, and 41. The g_2 coefficients refer to the splines in the average tenure with knots at one, two, three, and four years of tenure.

and learning-by-doing. The RW parameter σ_q^2 captures the effect of tenure on the permanent earnings component, and yields a marginally significant coefficient estimate, implying that persistent earnings differentials increase within job spells on top of life-cycle effects. We stress that the presence of tenure in the permanent component ensures that the results on tenure and instability, discussed later in this section, do not reflect the omission of tenure effects in the permanent component.[16]

[16] We further assessed that the results are robust to the specification of permanent earnings in two ways. First, we excluded tenure effects from the permanent component. Second, we specified tenure effects in the permanent component using the same exponential spline that we used in the transitory component. In both cases, notwithstanding the different modeling of tenure in the permanent component, the results on the effects of tenure on earnings instability were remarkably robust (see Cappellari and Leonardi, 2013).

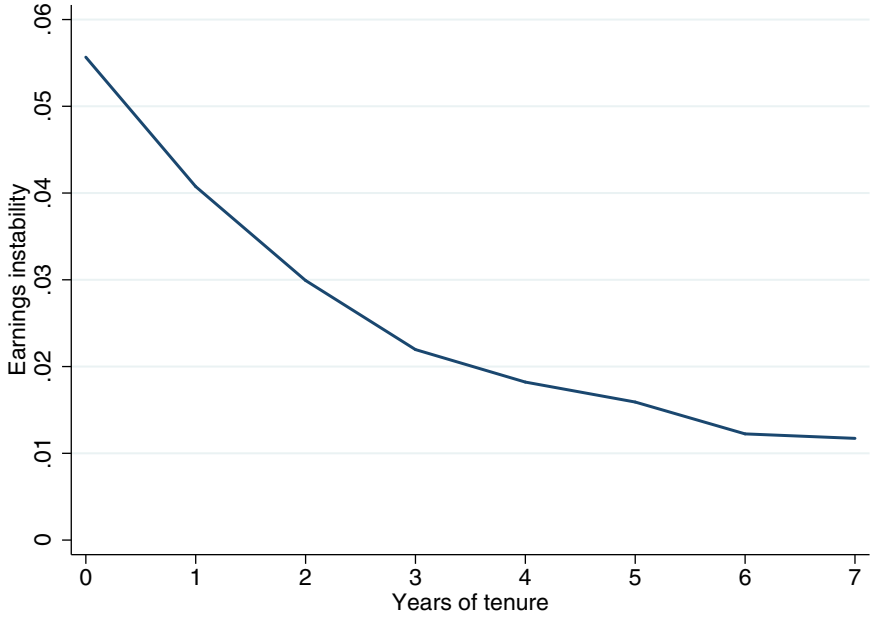

Fig. 3. Estimated earnings instability by tenure

The main parameters of interest in Table 4 are the g_2 coefficients that relate earnings instability with tenure (the spline has knots at one, two, three, and four years of tenure). Their estimates show that instability decreases with seniority on the job: tenure decreases rapidly over the first three years of the match, and then flattens out over the fourth year and afterwards. Importantly, the tenure effects that we estimate are obtained while controlling for the relation between earnings instability and age (through the g_1 coefficients), so that the result is net of any spurious influence that might emerge in the presence of a correlation between age and tenure. Other estimates of the AR(1) parameter in Column 1 of Table 4 reveal an intermediate degree of correlation of the shocks ($\rho = 0.42$), somewhat smaller than the estimate reported by Baker and Solon (2003) (i.e., $\rho = 0.54$), and a U-shaped pattern of instability between the mid-30s and the mid-50s, which is also consistent with the evidence in Baker and Solon (2003) of U-shaped life-cycle patterns of instability.

The predictions from this model are summarized in Figure 3, which plots the estimated earnings instability (the variance of transitory shocks σ_ε^2) against tenure. The predictions are averaged over cohorts. The figure shows a clear downward trend with tenure. More specifically, the average instability is 0.055 at the start of the job match, and 0.011 after seven years of tenure, implying a yearly reduction rate of approximately 11 percent

$(= (0.055 - 0.011)/(0.055 \times 7) \times 100)$. This reduction is concentrated in the first three years of the match, where the average yearly reduction rate is 20 percent $(= (0.055 - 0.022)/(0.055 \times 3) \times 100)$. This finding is consistent with the results of Lange (2007), who shows that much of the employer's learning occurs within the first three years of the match.

Results of the Model with Temporary Contracts

As an alternative way to test the idea that shorter tenure is associated with earnings instability, we now show estimates of earnings variance components that are parametrized with respect to the type of job contracts, either temporary or open-ended. The results from this exercise are presented in Column 2 of Table 4. The coefficients linking the contract type to permanent and transitory earnings shocks (γ and ϕ, respectively) attract the signs we would expect *a priori*, indicating that individuals on temporary contracts have, on average, a lower permanent variance of earnings and a higher instability than permanent workers. This lower permanent variance reflects a compressed distribution of long-term earnings for temporary workers, which can emerge insofar as temporary contracts are less favorable to the development of job-specific skills and are characterized by less training than open-ended contracts.[17] Using parameter estimates, we can predict the transitory earnings variance associated with temporary contracts; see Figure 4, where we average the predicted transitory variances over cells defined by the incidence of temporary contracts (i.e., below 5 percent and above 10 percent), and plot the estimated averages over time. There is a clear gap in the average transitory earnings variance between the cohorts with low and high incidences of temporary contracts. This gap is of the order of 70 percent and (with little variation) it is stable until 1998. After the introduction of the Treu reform, which liberalized temporary employment contracts, the gap rapidly rises until it reaches almost a 100 percent difference in 2003, the last year of data in our sample.

Results of the Model with Tenure Using Variation in Temporary Employment

The evidence from Figure 4 suggests that the reforms of temporary employment had an impact on earnings instability, and we argue that this

[17] We assessed the sensitivity of these findings on instability and the type of employment contract by estimating a version of the model in which contract types are not allowed to affect long-term earnings. The estimated ϕ coefficients on temporary contracts were very similar to the ones in Column 2 of Table 4, pointing to the robustness of the findings on temporary contracts and instability (see Cappellari and Leonardi, 2013).

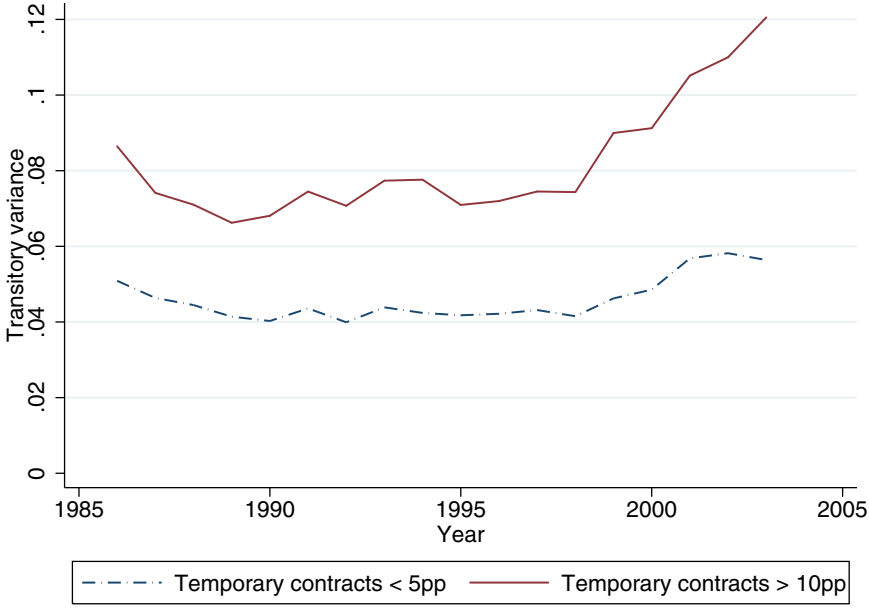

Fig. 4. Estimated transitory variance by incidence of temporary contracts

occurred through a reduction in tenure associated with the spread of temporary employment. We now pursue this idea more explicitly by taking into account the relation between temporary employment and tenure within the analytical framework of our model. We use equation (9) and estimate the impact of the cohort-specific incidence of temporary employment on individual workplace tenure while controlling for cohort and time fixed effects. The coefficient relating temporary employment to tenure (δ_F) is estimated to be equal to -4.33 and statistically significant (robust t-ratio $= 2.70$), indicating that an increase in the incidence of temporary employment from 0 to 100 reduces cohort members' tenure by 4.3 months, net of cohort and time fixed effects.[18] The estimated effect corresponds to an 8 percent reduction compared with the sample average (55 months of tenure).[19]

[18] The coefficients' estimates on cohort and time fixed effects are not reported, but are available upon request; they indicate that tenure declines with year of birth and increases over time.

[19] To illustrate the variation used to estimate δ_F, we can compare the estimates with those from a regression that ignores both time and cohort fixed effects (i.e., ignores between-cohort selectivity). This latter regression yields an estimate of δ_F equal to -110.23 (t-ratio $=$ 15.85), which suggests that there is a lot of selectivity between cohorts and time that is controlled for in equation (9).

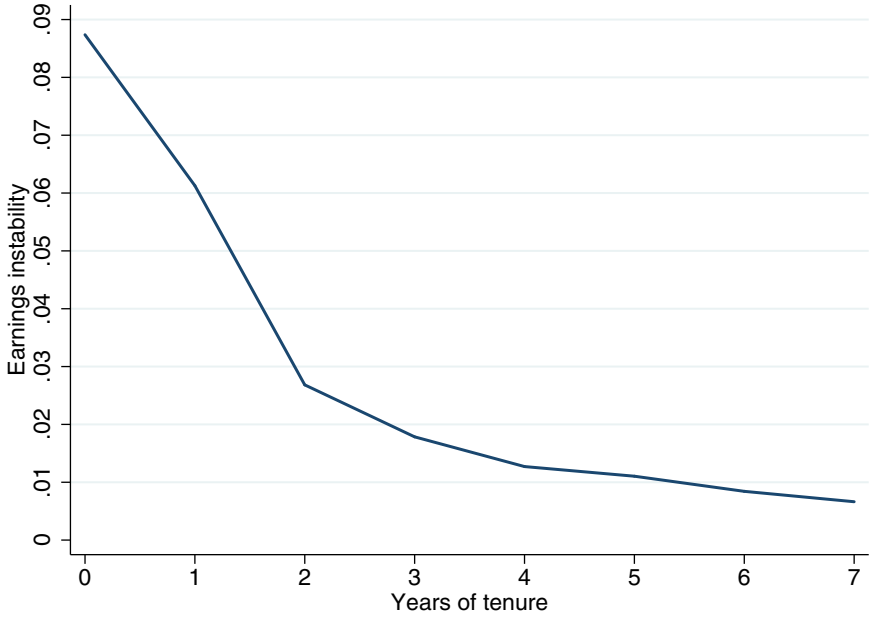

Fig. 5. Estimated earnings instability by tenure, model with predicted tenure

We use predicted tenure from equation (9) in place of actual tenure in our model of earnings instability, using weights that are proportional to the inverse of the prediction variance. The prediction varies both because of time and cohort fixed effects, and because of variation in the incidence of temporary employment on top of those effects. Because the earnings dynamics model also features time and cohort fixed effects, it is only the latter source of variation that provides identification of the effect of tenure on instability. The results from the model with predicted tenure are presented in Column 3 of Table 4. The parameter estimates on the permanent component are stable in comparison with their counterparts in Column 1. There is instead some difference in the parameter estimates for the transitory component, pointing to a greater earnings instability at the start of the job match, and a sharper decline in subsequent years, than in Column 1.

The overall pattern of instability over tenure is shown in Figure 5, which qualitatively confirms the pattern of Figure 3 and quantitatively shows a larger instability at the start of the match (0.0875 instead 0.055), followed by a steeper decline afterwards. In the first three years, the fall is by 26.5 percent on a yearly basis ($= (0.0875 - 0.018)/(0.0875 \times 3) \times 100$), which compares to the 20 percent of the model underlying Figure 3.

After seven years of tenure, earnings instability is similar to that of Figure 3; the overall average yearly decline of tenure is 13 percent ($= (0.0875 - 0.008)/(0.0875 \times 7) \times 100$). Because the model in Column 3 uses variation in tenure induced by a change in the type of employment contract, these results suggest that workers on short tenure because of the temporary employment contracts reforms have a level of instability that is higher than that experienced by permanent employees with equal tenure.

VI. Discussion and Conclusions

In this paper, we use Italian panel data to estimate the impact of on-the-job tenure on earnings instability. Although other papers (Huff Stevens, 2001; Hospido, 2012; Leonardi, 2012) have looked into the effect of voluntary and involuntary job changes on instability, we are the first to develop a formal model that takes into account tenure in the decomposition of the earnings variance. We find 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 an 11 percent reduction in instability. The reduction is even larger (13 percent) when we use the temporary employment contracts reforms to predict tenure.

Although this quantification of the effect of tenure on instability does not imply any particular interpretation, these results are potentially consistent with different models of wage determination. The results are consistent with matching models where overall earnings profiles tend to their long-term component as individuals settle down in their job and information on their ability is revealed. Lange (2007) finds that, in the US, the initial expectation error about match quality declines by 50 percent in three years, which approximately equals our estimate of a 20 percent yearly reduction in earnings instability in the first three years of the match (which becomes 26.5 percent in the model that uses predicted 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-contract setting is consistent with our results if the provision of insurance increases with tenure and leads to a decline of earnings instability.

We also look explicitly at the effect of temporary employment contracts on earnings instability. We find that, on average, cohorts of workers with a high incidence of temporary contracts experience a transitory earnings variance that is greater, by between 50 and 100 percent, than that for cohorts with a low incidence. The exercise of this paper is potentially

relevant from the policy point of view for many European countries, 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 are additional channels through which these new type of jobs affect individual welfare, namely through an increased uncertainty surrounding long-term earnings profiles.

Appendix: Additional Tables

Table A1. *Model with tenure: time shifters of permanent and transitory variance (1986 = 1)*

	Coeff.	SE		Coeff.	SE
π_{1987}	1.0377	0.0144	τ_{1987}	0.8978	0.0756
π_{1988}	1.0500	0.0163	τ_{1988}	0.9256	0.1001
π_{1989}	1.0379	0.0170	τ_{1989}	0.9501	0.1193
π_{1990}	1.0507	0.0181	τ_{1990}	0.9714	0.1337
π_{1991}	1.0305	0.0187	τ_{1991}	1.0712	0.1549
π_{1992}	1.0265	0.0201	τ_{1992}	1.0473	0.1595
π_{1993}	1.0102	0.0203	τ_{1993}	1.1154	0.1668
π_{1994}	0.9916	0.0212	τ_{1994}	1.1354	0.1782
π_{1995}	0.9985	0.0224	τ_{1995}	1.1392	0.1827
π_{1996}	0.9569	0.0223	τ_{1996}	1.1898	0.1949
π_{1997}	0.9459	0.0228	τ_{1997}	1.2479	0.2083
π_{1998}	0.9397	0.0234	τ_{1998}	1.2666	0.2131
π_{1999}	0.9358	0.0240	τ_{1999}	1.4437	0.2443
π_{2000}	0.9194	0.0237	τ_{2000}	1.4880	0.2549
π_{2001}	0.8853	0.0231	τ_{2001}	1.6787	0.2937
π_{2002}	0.8772	0.0235	τ_{2002}	1.7434	0.3073
π_{2003}	0.8481	0.0234	τ_{2003}	1.8530	0.3323

Table A2. *Model with tenure: cohort shifters and cohort initial conditions (1957 = 1)*

	Coeff.	SE		Coeff.	SE
λ_{1940}	0.6863	0.0303	η_{1940}	0.8919	0.3512
λ_{1941}	0.6628	0.0302	η_{1941}	0.3124	0.2129
λ_{1942}	0.7334	0.0311	η_{1942}	0.2410	0.2499
λ_{1943}	0.7396	0.0312	η_{1943}	0.4347	0.2986
λ_{1944}	0.7375	0.0291	η_{1944}	0.2484	0.1628
λ_{1945}	0.8001	0.0302	η_{1945}	0.2484	0.1628

(*Continued*)

Table A2. *Continued*

	Coeff.	SE		Coeff.	SE
λ_{1946}	0.8065	0.0291	η_{1946}	0.8272	0.3276
λ_{1947}	0.7691	0.0291	η_{1947}	0.4202	0.2163
λ_{1948}	0.8030	0.0280	η_{1948}	0.8527	0.3151
λ_{1949}	0.7981	0.0282	η_{1949}	1.0298	0.2994
λ_{1950}	0.8124	0.0264	η_{1950}	0.9823	0.3234
λ_{1951}	0.8250	0.0281	η_{1951}	1.2478	0.5071
λ_{1952}	0.8908	0.0306	η_{1952}	0.7415	0.2620
λ_{1953}	0.9095	0.0305	η_{1953}	0.7803	0.2483
λ_{1954}	0.9123	0.0329	η_{1954}	0.6366	0.2372
λ_{1955}	0.9679	0.0330	η_{1955}	1.0981	0.3389
λ_{1956}	1.0195	0.0328	η_{1956}	1.4030	0.3985
λ_{1958}	1.0743	0.0347	η_{1958}	0.9566	0.2841
λ_{1959}	1.1064	0.0344	η_{1959}	1.0359	0.2758
λ_{1960}	1.1164	0.0362	η_{1960}	1.0164	0.2792
λ_{1961}	1.0901	0.0342	η_{1961}	1.2588	0.3415
λ_{1962}	1.1498	0.0377	η_{1962}	1.1675	0.3134
λ_{1963}	1.1776	0.0370	η_{1963}	1.3312	0.3347
λ_{1964}	1.1329	0.0368	η_{1964}	1.2100	0.2925
λ_{1965}	1.2259	0.0412	η_{1965}	1.2852	0.3382
λ_{1966}	1.2319	0.0416	η_{1966}	1.1969	0.3718
λ_{1967}	1.2728	0.0453	η_{1967}	0.8725	0.3096
λ_{1968}	1.2757	0.0464	η_{1968}	0.7129	0.2731
λ_{1969}	1.2914	0.0495	η_{1969}	1.1217	0.4647
λ_{1970}	1.3304	0.0567	η_{1970}	0.8444	0.3465
λ_{1971}	1.4366	0.0608	η_{1971}	0.8953	0.3925
λ_{1972}	1.3654	0.0660	η_{1972}	0.8548	0.3887
λ_{1973}	1.3198	0.0680	η_{1973}	1.0897	0.4726

Table A3. *Model with temporary contracts: time shifters of permanent and transitory variance (1986 = 1)*

	Coeff.	SE		Coeff.	SE
π_{1987}	1.0401	0.0147	τ_{1987}	0.8583	0.0454
π_{1988}	1.0523	0.0164	τ_{1988}	0.8192	0.0519
π_{1989}	1.0403	0.0171	τ_{1989}	0.7882	0.0522
π_{1990}	1.0518	0.0181	τ_{1990}	0.7794	0.0521
π_{1991}	1.0283	0.0187	τ_{1991}	0.8135	0.0529
π_{1992}	1.0239	0.0199	τ_{1992}	0.7800	0.0504
π_{1993}	1.0080	0.0207	τ_{1993}	0.8179	0.0526
π_{1994}	0.9915	0.0213	τ_{1994}	0.8026	0.0516
π_{1995}	0.9984	0.0224	τ_{1995}	0.7937	0.0512
π_{1996}	0.9542	0.0222	τ_{1996}	0.8042	0.0519
π_{1997}	0.9414	0.0225	τ_{1997}	0.8204	0.0533
π_{1998}	0.9329	0.0230	τ_{1998}	0.8205	0.0535
π_{1999}	0.9237	0.0237	τ_{1999}	0.8984	0.0589
π_{2000}	0.9045	0.0235	τ_{2000}	0.9017	0.0599
π_{2001}	0.8667	0.0229	τ_{2001}	0.9690	0.0644
π_{2002}	0.8567	0.0233	τ_{2002}	0.9912	0.0661
π_{2003}	0.8258	0.0232	τ_{2003}	1.0342	0.0699

Table A4. *Model with temporary contracts: cohort shifters and cohort initial conditions (1957 = 1)*

	Coeff.	SE		Coeff.	SE
λ_{1940}	0.6638	0.0301	η_{1940}	0.7234	0.2604
λ_{1941}	0.6416	0.0301	η_{1941}	0.3002	0.1779
λ_{1942}	0.7133	0.0313	η_{1942}	0.2223	0.2026
λ_{1943}	0.7199	0.0312	η_{1943}	0.4052	0.2488
λ_{1944}	0.7189	0.0293	η_{1944}	0.2574	0.1350
λ_{1945}	0.7821	0.0306	η_{1945}	0.2574	0.1350
λ_{1946}	0.7890	0.0293	η_{1946}	0.7577	0.2687
λ_{1947}	0.7532	0.0292	η_{1947}	0.3741	0.1709
λ_{1948}	0.7878	0.0281	η_{1948}	0.7600	0.2396
λ_{1949}	0.7837	0.0285	η_{1949}	0.9210	0.2360
λ_{1950}	0.7988	0.0267	η_{1950}	0.9097	0.2549
λ_{1951}	0.8129	0.0286	η_{1951}	1.0462	0.3685
λ_{1952}	0.8815	0.0315	η_{1952}	0.7069	0.2080
λ_{1953}	0.9006	0.0313	η_{1953}	0.8116	0.2124
λ_{1954}	0.9056	0.0339	η_{1954}	0.6519	0.2013
λ_{1955}	0.9634	0.0342	η_{1955}	1.0170	0.2620
λ_{1956}	1.0179	0.0341	η_{1956}	1.2897	0.3079
λ_{1958}	1.0768	0.0364	η_{1958}	1.0071	0.2377
λ_{1959}	1.1111	0.0363	η_{1959}	1.0723	0.2301
λ_{1960}	1.1248	0.0385	η_{1960}	1.0248	0.2269
λ_{1961}	1.0976	0.0367	η_{1961}	1.1544	0.2580
λ_{1962}	1.1596	0.0407	η_{1962}	1.1492	0.2553
λ_{1963}	1.1935	0.0402	η_{1963}	1.2656	0.2632
λ_{1964}	1.1431	0.0404	η_{1964}	1.1042	0.2237
λ_{1965}	1.2495	0.0454	η_{1965}	1.1828	0.2544
λ_{1966}	1.2554	0.0467	η_{1966}	1.2339	0.2901
λ_{1967}	1.3041	0.0509	η_{1967}	1.2848	0.3158
λ_{1968}	1.3022	0.0528	η_{1968}	1.2253	0.3035
λ_{1969}	1.3129	0.0565	η_{1969}	1.8036	0.4791
λ_{1970}	1.3672	0.0659	η_{1970}	1.6517	0.4115
λ_{1971}	1.5052	0.0712	η_{1971}	1.8316	0.4577
λ_{1972}	1.4222	0.0793	η_{1972}	1.7331	0.4771
λ_{1973}	1.3578	0.0832	η_{1973}	2.0725	0.5356

Table A5. *Model with estimated tenure: time shifters of permanent and transitory variance (1986 = 1)*

	Coeff.	SE		Coeff.	SE
π_{1987}	1.0394	0.0144	τ_{1987}	0.9372	0.0847
π_{1988}	1.0508	0.0161	τ_{1988}	1.0314	0.1153
π_{1989}	1.0364	0.0168	τ_{1989}	1.1234	0.1455
π_{1990}	1.0483	0.0177	τ_{1990}	1.1815	0.1649
π_{1991}	1.0276	0.0182	τ_{1991}	1.3354	0.1957
π_{1992}	1.0226	0.0194	τ_{1992}	1.3282	0.2051
π_{1993}	1.0095	0.0200	τ_{1993}	1.3741	0.2024
π_{1994}	0.9900	0.0206	τ_{1994}	1.3974	0.2158
π_{1995}	0.9965	0.0216	τ_{1995}	1.4071	0.2234
π_{1996}	0.9545	0.0215	τ_{1996}	1.4981	0.2466

(*Continued*)

Table A5. *Continued*

	Coeff.	SE		Coeff.	SE
π_{1997}	0.9428	0.0218	τ_{1997}	1.5973	0.2715
π_{1998}	0.9364	0.0224	τ_{1998}	1.6356	0.2818
π_{1999}	0.9334	0.0232	τ_{1999}	1.8704	0.3256
π_{2000}	0.9179	0.0230	τ_{2000}	1.9288	0.3406
π_{2001}	0.8849	0.0226	τ_{2001}	2.1892	0.3960
π_{2002}	0.8771	0.0231	τ_{2002}	2.2842	0.4173
π_{2003}	0.8487	0.0231	τ_{2003}	2.4348	0.4542

Table A6. *Model with estimated tenure: cohort shifters and cohort initial conditions (1957 = 1)*

	Coeff.	SE		Coeff.	SE
λ_{1940}	0.6898	0.0304	η_{1940}	0.8955	0.3559
λ_{1941}	0.6661	0.0301	η_{1941}	0.3024	0.2125
λ_{1942}	0.7368	0.0312	η_{1942}	0.2298	0.2508
λ_{1943}	0.7430	0.0313	η_{1943}	0.4242	0.2985
λ_{1944}	0.7407	0.0291	η_{1944}	0.2373	0.1623
λ_{1945}	0.8033	0.0302	η_{1945}	0.2373	0.1623
λ_{1946}	0.8095	0.0290	η_{1946}	0.8181	0.3274
λ_{1947}	0.7719	0.0290	η_{1947}	0.4200	0.2182
λ_{1948}	0.8058	0.0277	η_{1948}	0.8486	0.3187
λ_{1949}	0.8006	0.0281	η_{1949}	1.0277	0.3011
λ_{1950}	0.8146	0.0262	η_{1950}	0.9777	0.3263
λ_{1951}	0.8269	0.0279	η_{1951}	1.2649	0.5176
λ_{1952}	0.8923	0.0304	η_{1952}	0.7446	0.2654
λ_{1953}	0.9106	0.0302	η_{1953}	0.7658	0.2473
λ_{1954}	0.9133	0.0326	η_{1954}	0.6318	0.2380
λ_{1955}	0.9683	0.0327	η_{1955}	1.1012	0.3428
λ_{1956}	1.0195	0.0325	η_{1956}	1.4112	0.4034
λ_{1958}	1.0735	0.0345	η_{1958}	0.9463	0.2846
λ_{1959}	1.1052	0.0342	η_{1959}	1.0221	0.2755
λ_{1960}	1.1146	0.0359	η_{1960}	1.0051	0.2795
λ_{1961}	1.0882	0.0340	η_{1961}	1.2648	0.3463
λ_{1962}	1.1474	0.0374	η_{1962}	1.1599	0.3119
λ_{1963}	1.1741	0.0368	η_{1963}	1.3391	0.3391
λ_{1964}	1.1299	0.0367	η_{1964}	1.2267	0.2978
λ_{1965}	1.2211	0.0413	η_{1965}	1.3037	0.3455
λ_{1966}	1.2273	0.0416	η_{1966}	1.1035	0.3531
λ_{1967}	1.2672	0.0451	η_{1967}	0.6911	0.2515
λ_{1968}	1.2706	0.0462	η_{1968}	0.5073	0.1976
λ_{1969}	1.2873	0.0492	η_{1969}	0.7595	0.3136
λ_{1970}	1.3251	0.0563	η_{1970}	0.5333	0.2169
λ_{1971}	1.4302	0.0606	η_{1971}	0.5439	0.2373
λ_{1972}	1.3652	0.0654	η_{1972}	0.5736	0.2497
λ_{1973}	1.3228	0.0669	η_{1973}	0.7282	0.3060

References

Adda, J., Dustmann, C., Meghir, C., and Robin, J. M. (2013), Career Progression, Economic Downturns and Skills, NBER Working Paper 18832.

Alessie, R. and Kalwij, A. (2007), Permanent and Transitory Wage Inequality of British Men, 1975–2001: Year, Age and Cohort Effects, *Journal of Applied Econometrics 22*, 1033–1062.

Altonji, J. G. and Williams, N. (2005), Do Wages Rise with Job Seniority? A Reassesment, *Industrial and Labor Relations Review 58*, 370–397.

Altonji, J. G., Smith, A., and Vidangos, I. (2013), Modeling Earnings Dynamics, *Econometrica 81*, 1395–1454.

Autor, D. H. and Houseman, S. N. (2010), Do Temporary-Help Jobs Improve Labor Market Outcomes for Low-Skilled Workers? Evidence from Work First, *American Economic Journal: Applied Economics 2*, 96–128.

Baker, M. and Solon, G. (2003), Earnings Dynamics and Inequality among Canadian Men 1976–1992, *Journal of Labor Economics 21*, 289–321.

Ben-Porath, Y. (1967), The Production of Human Capital and the Life Cycle of Earnings, *Journal of Political Economy 75*, 352–365.

Bingley, P., Cappellari, L., and Westergård-Nielsen, N. (2013), Unemployment Insurance, Wage Dynamics and Inequality over the Life Cycle, *Economic Journal 123*, 341–372.

Blundell, R., Graber, M., and Mogstad, M. (2015), Labor Income Dynamics and the Insurance from Taxes, Transfers and the Family, *Journal of Public Economics 127*, 58–73.

Booth, A., Francesconi, M., and Frank, J. (2002), Temporary Jobs: Stepping Stones or Dead Ends?, *Economic Journal 112*, F585–F606.

Brandolini, A., Cipollone, P., and Sestito, P. (2002), Earnings Dispersion, Low Pay and Household Poverty in Italy, 1977–1998 in D. Cohen, T. Piketty, and G. Saint-Paul (eds.), *The Economics of Rising Inequalities*, Oxford University Press, Oxford, 225–264.

Cameron, S. and Tracy, J. (1998), Earnings Variability in the United States: An Examination Using Matched-CPS Data (available at http://www.newyorkfed.org/research/economists/tracy/earnings_variability.pdf).

Cappellari, L. (2002), Earnings Dynamics and Uncertainty in Italy: How Do They Differ between the Private and Public Sectors?, *Labour Economics 9*, 477–496.

Cappellari, L. (2004), The Dynamics and Inequality of Italian Men's Earnings: Long Term Changes or Transitory Fluctuations?, *Journal of Human Resources 2*, 475–499.

Cappellari, L. and Leonardi, M. (2013), Earnings Instability and Tenure, CESIfo working paper 4145.

Cappellari, L. and Jenkins, S. P. (2014), Earnings and Labour Market Volatility in Britain, with a Transatlantic Comparison, *Labour Economics 30*, 201–211.

Celik, S., Juhn, C., McCue, K., and Thompson, J. (2012), Understanding Earnings Instability: How Important are Employment Fluctuations and Job Changes, *B.E. Journal of Economic Analysis and Policy 12*, 1–27.

Cunha, F. and Heckman, J. (2007), The Evolution of Inequality, Heterogeneity and Uncertainty in Labor Earnings in the U.S. Economy, NBER Working Paper 13526.

Cunha, F., Heckman, J., and Navarro, S. (2005), Separating Uncertainty from Heterogeneity in Life Cycle Earnings, *Oxford Economic Papers 57*, 191–261.

Dahl, M., DeLeire, T., and Schwabish, J. (2011), Estimates of Year-to-Year Volatility in Earnings and in Household Incomes from Administrative, Survey, and Matched Data, *Journal of Human Resources 46*, 750–774.

Dickens, R. (2000), The Evolution of Individual Male Earnings in Great Britain: 1975–95, *Economic Journal 110*, 27–49.

Dustmann, C. and Meghir, C. (2005), Wages, Experience and Seniority, *Review of Economic Studies 72*, 77–108.
Dynan, K. E., Elmendorf, D. W., and Sichel, D. E. (2012), The Evolution of Household Income Volatility, *B.E. Journal of Economic Analysis and Policy 12*, Advances Article 3.
Farber, H. S. (2010), Job Loss in the Great Recession: Historical Perspective from the Displaced Workers Survey, 1984–2010, NBER Working Paper 17040.
Flabbi, L. and Leonardi, M. (2010), Sources of Earnings Instability: Estimates from an On-the-Job Search Model of the U.S. Labor Market, *European Economic Review 54*, 832–854.
Gladden, T. and Taber, C. (2009), The Relationship between Wage Growth and Wage Levels, *Journal of Applied Econometrics 24*, 914–932.
Gottschalk, P. and Moffitt, R. (1994), The Growth of Earnings Instability in the U.S. Labor Market, *Brookings Papers on Economic Activity 2*, 217–254.
Gottschalk, P. and Moffitt, R. (1999), Changes in Job Instability and Insecurity Using Monthly Survey Data, *Journal of Labor Economics 17*, S91–S126.
Gottschalk, P. and Moffitt, R. (2009), The Rising Instability of U.S. Earnings, *Journal of Economic Perspectives 23*, 3–24.
Guiso, L., Pistaferri, L., and Schivardi, F. (2005), Insurance within the Firm, *Journal of Political Economy 113*, 1054–1085.
Guvenen, F. (2007), Learning Your Earning: Are Labor Income Shocks Really Very Persistent?, *American Economic Review 97*(3), 687–712.
Haider, S. (2001), Earnings Instability and Earnings Inequality of Males in the U.S.: 1967–1991, *Journal of Labor Economics 19*, 799–836.
Heathcote, J., Storesletten, K., and Violante, G. (2010), The Macroeconomic Implications of Rising Wage Inequality in the United States, *Journal of Political Economy 118*, 681–722.
Heathcote, J., Storesletten, K., and Violante, G. (2014), Consumption and Labor Supply with Partial Insurance: An Analytical Framework, *American Economic Review 104*(7), 2075–2126.
Hospido, L. (2010), Wage Dynamics in the Presence of Unobserved Individual and Job Heterogeneity, IZA Discussion Paper 5088.
Hryshko, D. (2012), Labor Income Profiles are not Heterogeneous: Evidence from Income Growth Rates, *Quantitative Economics 3*, 177–209.
Huff Stevens, A. (2001), Changes in Earnings Instability and Job Loss, *Industrial and Labor Relations Review 55*, 60–78.
Jaeger, D. A. and Huff Stevens, A. (1999), Is Job Stability in the United States Falling? Reconciling Trends in the Current Population Survey and Panel Study of Income Dynamics, *Journal of Labor Economics 17*, S1–S28.
Lange, F. (2007), The Speed of Employer Learning, *Journal of Labor Economics 25*, 1–35.
Leonardi, M. (2012), Job Mobility and Earnings Instability, University of Milan.
Low, H., Meghir, C., and Pistaferri, L. (2010), Wage Risk and Employment Risk over the Life Cycle, *American Economic Review 100*(4), 1432–1467.
Meghir, C. and Pistaferri, L. (2004), Income Variance Dynamics and Heterogeneity, *Econometrica 72*, 1–32.
Meghir, C. and Pistaferri, L. (2011), Earnings, Consumption and Life Cycle Choices, in O. Ashenfelter and D. Card (eds.), *Handbook of Labor Economics*, Volume 4B, Elsevier, Amsterdam, 773–854.
Mincer, J. (1958), Investment in Human Capital and Personal Income Distribution, *Journal of Political Economy 65*, 281–302.
Moffitt, R. and Gottschalk, P. (2012), Trends in the Transitory Variance of Male Earnings in the U.S., 1970–2004, *Journal of Human Resources 47*, 204–236.
Parent, D. (2002), Matching, Human Capital and the Covariance Structure of Earnings, *Labour Economics 9*, 375–404.

Ramos, X. (2003), The Covariance Structure of Earnings in Great Britain 1991–1999, *Economica 70*, 353–374.
Shin, D. and Solon, G. (2011), Trends in Men's Earnings Volatility: What Does the Panel Study of Income Dynamics Show?, *Journal of Public Economics 95*, 973–982.
Venn, D. (2011), Earnings Volatility and Its Consequences for Households, OECD Social, Employment and Migration Working Paper 125.
Ziliak, J., Hardy, B., and Bollinger, C. (2011), Earnings and Income Volatility in America: Evidence from Matched CPS, *Labour Economics 18*, 742–754.

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