THE ROLE OF WAGE DIFFERENCES AND INDIVIDUAL LABOUR SUPPLY ON MALE EARNINGS INEQUALITY: EMPIRICAL EVIDENCE FROM SPAIN[*]

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ABSTRACT

This paper studies the link between wage differences and individual earnings inequality for a sample of Spanish continuously working males from the European Community Household Panel for the period from 1994 to 1998. We analyse and quantify the contribution of two labour market features that could affect this link: the role of permanent and transitory wage inequality, and the existence of a significant response in individuals' hours of work to shifts in their own wage rates. Based on Hyslop (2001), we propose a particular specification for wages and earnings that incorporates an intertemporal labour supply model and we obtain predictions on wage and earnings inequality using a covariance structure framework.

Keywords: Wage inequality, individual earnings inequality, individual labour supply, covariance structure model.

JEL: C23, J22.

RESUMEN

En este trabajo se analiza la relación entre las diferencias salariales y la desigualdad en rentas laborales individuales para una muestra de hombres empleados de forma continua en España en el período comprendido entre 1994 y 1998 procedente del Panel de Hogares de la Unión Europea (PHOGUE). En concreto, se analiza y cuantifica la contribución de dos aspectos del mercado laboral que pueden afectar a dicha relación o transmisión: el papel de la desigualdad salarial permanente o transitoria, y la existencia de respuestas significativas de las horas de trabajo de los individuos a variaciones en sus propios salarios. Basándonos en Hyslop (2001), se propone una especificación particular para los salarios y las rentas laborales que incorpora un modelo de oferta de trabajo intertemporal. A través de la estimación de dicho modelo, se obtienen predicciones de la desigualdad salarial y la desigualdad en rentas laborales utilizando un enfoque de estructura de covarianza.

Palabras clave: Desigualdad salarial, desigualdad en rentas laborales, oferta de trabajo individual, modelo de estructura de covarianza.

1 Introduction

The issue of the changes of the wage structure and overall earnings inequality has been extensively studied in Labour Economics. The normative interest in wage inequality concerns its effect on the individual well-being distribution since widening wage structure could imply widening earnings and consumption inequality, and associated welfare problems. In the literature, there is a very important and wide research on the changes in the wage structure and earnings inequality for the United States, specially for the last two decades (Abowd and Card (1989), Bound and Johnson (1992), Katz and Murphy (1992), Hyslop (2001)). A great part of this literature analyses the evolution of wage differences between and within different groups of workers and measures the contributions of factors as trade, human capital and technology on these differences (Freeman (1984),Willis (1986), Katz and Murphy (1992), Bound and Johnson (1992), Autor et al (1998)). The reason for this common interest is basically twofold. First, the wage structure has changed significantly in most developed countries in last decades. Even though wage inequality decreased substantially during the 70's, a great expansion of wage differentials arised in the 80's, specially in the United States and United Kingdom (see Freeman and Katz (1994), (1995)). Second, an increasing number of appropriate and comparable micro datasets recently have been available for performing these studies during the last decades. However, as Hyslop (2001) notes there is a lack of research on the link between wage differences and individual earnings inequality and the mechanisms underlying this link in spite of the importance of this question from a policy point of view.[1]

For Spain, as Jimeno et al. (2001) point out, the evidence on the evolution of wage and earnings inequality is rather scarce. Most of the studies has focused on the dispersion in wages across workers (García-Perea (1991), Jimeno and Toharia (1994)), and the

[1] Specifically, Hyslop (2001) analyse the link between the dispersion in rates of pay and family earnings inequality.

contributions of different economic individual charateristics to the wage inequality (Bover et al. (2000) and Jimeno et al. (2001)) using different datasets. More recently, Cervini and Ramos (2006) analyse the evolution of earnings inequality for the period from 1993 to 2000 for a sample of male full-time employees drawn from the European Community Household Panel (ECHP). Specifically, they use an error component model to decompose the earnings covariance structure into its permanent and transitory components. They show that cross-sectional earnings inequality for this sample has not changed substantially over the period. In addition to this, they find that this small change is determined by a rise in the permanent component of the variance and a decrease in the transitory component. In particular, they claim that these results could be explained by the labour market reforms implemented in Spain in 1994 and 1997, which introduced more stability to temporary employment contracts and pushed towards a decrease in temporary employment. However, Cervini and Ramos (2006) do not analyse the connection between wage and earnings inequality. Therefore, given the policy implications of wage and earnings inequality in terms of consumption inequality and individuals' welfare, the goal of this paper is to document the evolution of these two elements for continuously working males in Spain for the period from 1994 to 1998. In particular, we analyse the link between individual wages dispersion and the inequality of the individual earnings and we perform a quantitative decomposition of the elements that affect this transmission. Following Hyslop (2001), we propose a model for wages and earnings that is based on an intertemporal individual labour supply model and we obtain predictions for wage and earnings inequality using a covariance structure framework. Specifically, we focus on two labour market features that could affect this link: first, the distinction between permanent and transitory shocks and their relative contributions on wage inequality. Second, the existence of significant responses in individuals' hours of work to shifts in their own wage rates. This behaviour could affect the transmission of wage differences to earnings inequality since annual earnings are defined as the product of wages and the individual's hours of work per year. In particular, if there is no labour supply responses to changes in hourly wages, wage inequality is transmitted perfectly to individual earnings inequality. However, the existence of a significant labour supply behaviour could alter this transmission through income and substitution effects. An income effect in which each individual's hours of work respond negatively to a variation in his wage will moderate the increase in earnings inequality. However, a substitution effect in which each individual's hours of work react positively to changes in his wage will emphasize the increase in inequality. These two aspects are studied for a sample of Spanish continuously working males from the ECHP for the period from 1994 to 1998.

This paper proceeds as follows. Section 2 presents a brief description of the data and the trends of wage and individual earnings cross-sectional inequality for our Spanish sample. In section 3, we specify a statistical model for the wages and earnings processes to obtain predictions for the relationship between wage and earnings inequality. Based on this theoretical model we develop an econometric specification for wages and we fit this model to the empirical covariance structure of wages. Next, we present a particular intertemporal individual labour supply model that incorporates the empirical specification for wages and reflects the relationship between wage and individual earnings processes. Section 4 presents the results for the estimation of the joint model of wages and earnings and section 5 focuses on the predictions implied by these estimates in terms of earnings inequality. Finally, section 6 concludes.

2 Data and Descriptive Analysis

2.1 Data

The data used in this analysis is drawn from the Spanish version of the ECHP. This data set provides comparable statistical annual information about the labour status and welfare level of households in the EU-15 countries allowing their social and economic situation to be analysed. For Spain, this is the unique panel that provides detailed information at household and individual level on earnings, income, hours of work, and demographics of people aged 16 and older. Furthermore, it also contains information about the individual's job like economic activity, firm size, public/private sector, and region. However, there are not details about institutional factors (unions, government wage regulation, internal labour markets) that may influence how wages are determined.[2]

[2] In Spain there exists evidence about the significant positive effect of the dimension of unions on wage differences for data from the Encuesta de Estructura Salarial 1995/2002 (Jimeno et al. (2001), Card and De la Rica (2006)).

In particular, we consider we consider males living in Spain for the period from 1994 to 1999. However, since annual income variables refer to the period prior to the interview and the remaining data refer to the current period, the last year is lost for the estimation. We select males aged between 25 and 65 and continuously working as employees at each year. In order to maximise our sample size, we construct an unbalanced panel for the period 1994-1998 including every individual-year observation with information about all the variables that we consider in the analysis for a given year. As a result, we end up with a sample of 2814 men with a total of 7548 individual-year observations.

We define individual earnings as the individual's annual net labour income,[3] that has been deflacted by the annual mean of the Consumer Price Index (CPI) base 1992 and is expressed in euros. Individual's hourly wage rate is computed as the ratio of annual earnings and the number of hours of work per year.[4]

2.2 Wage and Earnings Inequality Trends

In this section, we provide a descriptive analysis of the distribution of wages and earnings in our sample of males for the period from 1994 to 1998. In addition to this, we analyse the evolution of the cross-sectional wage and individual earnings inequality over these years.

Table 1 reports some descriptive statistics of the distribution of the log of wages and the log of earnings across individuals for each year. According to this information, the mean of both variables decreases for the first three years and then, it increases in 1997 and 1998. With respect to the evolution of the standard deviation, we can see that the dispersion in the distribution of both variables has not remained constant. In particular, it has decreased for the overall period even though it presents a small increase for earnings

[3] In our sample, labour income for workers represents on average 95 percent of total income. Therefore, the analysis of earnings inequality should be taken as an accurate indicator of income inequality in the working population.

[4] In order to check the quality and the validity of our data, we compared the distribution of weekly hours of work using two comparable samples from the ECHP and the Labour Force Survey (EPA), respectively, for the year 1999. We obtained that these distributions were quite similar. In addition, we did a similar comparison for the distribution of the hourly wage rate from the ECHP and the corresponding measure from the Encuesta de Estructura Salarial for the year 1995. We obtained very similar results for both surveys.

in 1995 and a larger increase for the year 1996 for both variables.

Since the sample means of both wages and earnings do not change substantially along these years, we will consider the cross-sectional variance as a measure of cross-sectional inequality.[5] Table 2 presents the trends for the cross-sectional inequality and its percentage changes during the period for both variables. We can observe that cross-sectional inequality is higher for wages than for earnings along the period since the variance for the log of wage is larger and the mean is smaller in every year. In addition, there has been a decrease in both wage and earnings cross-sectional inequality for the overall period although this reduction is more important for earnings. In fact, wage inequality has decreased almost 3 percent whereas earnings inequality has reduced almost 5 percent. However, if we consider each year variation, we can see that both variances present some year to year fluctuations. In particular, we can see that cross-sectional wage inequality decreases in all periods with the exception of 1996, in which it experiences an increase of 8 percent. Regarding earnings inequality, it increases in 1995 and 1996, and then it decreases during the rest of the period compensating the increase at the initial years.

To understand and identify the forces that are driven this kind of patterns, it seems crucial to raise three important questions. First, to determine to what extent changes in wage and earnings inequality are due to changes in the observable individuals' characteristics (between-group inequality) rather than unobservable characteristics (within-group or residual inequality). Second, to quantify the contributions of permanent and transitory components on wage and earnings inequality. This decomposition is of interest because it allows to determine the persistency of inequality changes and its potential effects on individuals' welfare since, in general, persistent differences affect more the distribution of well-being than transitory differences. Finally, to analyse the relative importance of the potential individual's labour supply behaviour on the evolution of earnings inequality. To quantify the role of all of these elements, we perform a standard variance decomposition for wages and earnings using the model specified in next sections.

[5] As it is pointed by Hyslop (2001), Juhn et al. (1993) and Gottschalk and Moffit (1994) the variance of the log of earnings has perhaps been the most commonly used measured of dispersion in the recent literature.

Table 1. Descriptive Statistics for Log Wage$^{(1)}$ and Log Earnings$^{(2)}$

		1994	1995	1996	1997	1998
	Mean	1.502	1.501	1.486	1.572	1.577
	Std. Dev	0.497	0.493	0.513	0.493	0.489
	Median	1.487	1.456	1.439	1.538	1.537
Log Wages	Min	-1.322	-3.128	-1.868	-1.903	-0.849
	25^{th} Quantile	1.212	1.209	1.182	1.238	1.258
	75^{th} Quantile	1.786	1.793	1.787	1.920	1.881
	Max	3.051	3.738	3.833	3.370	3.351
	Mean	9.236	9.224	9.206	9.272	9.279
	Std. Dev	0.473	0.476	0.494	0.462	0.461
	Median	9.183	9.166	9.148	9.198	9.225
Log Earnings	Min	6.436	4.917	6.035	6.223	7.197
	25^{th} Quantile	8.948	8.928	8.897	8.955	8.973
	75^{th} Quantile	9.487	9.481	9.456	9.575	9.564
	Max	10.914	11.938	11.878	11.233	11.194
	Sample Size	1456	1413	1338	1676	1665

Note: (1) Log of hourly wage.(2) Log of annual earnings. Wages and earnings have been deflacted by the annual mean of the Consumer Price Index (CPI) base 1992 and are expressed in euros.Number of observations, 1456 for 1994, 1413 for 1995, 1338 for 1996, 1676 for 1997 and 1665 for 1998.

Table 2. Cross-Sectional Variances, 1994-1998.

	Wage$^{(1)}$	Earnings$^{(2)}$
1994	0.2466	0.2236
1995	0.2432	0.2270
Percentage change (1994-1995)	(-1.4071)	(1.5050)
1996	0.2635	0.2445
Percentage change (1995-1996)	(8.3392)	(7.7058)
1997	0.2433	0.2137
Percentage change (1996-1997)	(-7.6522)	(-12.6017)
1998	0.2398	0.2128
Percentage change (1997-1998)	(-1.4698)	(-0.4193)
Percentage change (1994-1998)	(-2.8088)	(-4.8509)

Note: (1) Log of hourly wage. (2) Log of annual earnings. Wages and earnings have been deflacted by the annual mean of the Consumer Price Index (CPI) base 1992 and are expressed in euros. Number of observations, 1456 for 1994, 1413 for 1995, 1338 for 1996, 1676 for 1997 and 1665 for 1998.

3 An Empirical Model for Wages and Earnings

3.1 The Theoretical Specification

In this section, we propose a specific statistical model of covariance structure to perform a factor decomposition of wage and individual earnings inequality. In addition to this, we evaluate the elements that affect the link between wage differences and earnings inequality. Specifically, this model allows us to analyse the contribution of three different aspects: first, to what extent both cross-sectional variances are explained by some observable individuals' characteristics, and how much remained unexplained; second, the decomposition of both variables into permanent and transitory components; third, the role of individuals' labour supply decisions on the link between wage and earnings inequality.

We start with the following simple wage model:

$$\ln w_{it} = \Gamma_t + \psi_t^{'} Z_{it} + \theta_t \alpha_i + u_{it} \tag{1}$$

where the subscripts i denotes the individual, t denotes the year, and $\ln w_{it}$ is the log of real hourly wage of individual i in period t. The wage is decomposed into several terms: Γ_t represents the mean wage in period t and reflects common aggregate effects for all individuals, Z_{it} is a vector of observable individuals' characteristics, α_i is the unobserved permanent wage component that measures persistent worker skills such as ability and u_{it} represents the unobserved transitory wage component that captures the effects of labour market shocks. Both unobserved components, α_i and u_{it}, are assumed to be orthogonal with $E(\alpha_i) = E(u_{it}) = 0$, and with variances σ_α^2 and σ_u^2, respectively. The coefficients θ_t are time-varying factor loadings on the permanent wage component and can be interpreted as the time -varying skill price (returns to skill). By introducing different factor loadings for permanent wage components, we allow for nonstationarity in these components and different changes in permanent inequality.

From this theoretical model of wages, we can derive a theoretical specification for individual earnings since annual earnings are computed as the product of the hourly wage and the annual hours of work. Using this definition, we have that $\ln y_{it} = \ln w_{it} + \ln h_{it}$ where $\ln y_{it}$ and $\ln h_{it}$ represent the log of annual earnings and the log of annual hours or work of individual i in period t, respectively. This implies that the magnitude and the evolution of earnings differences are not only determined by the wage but also by the individual's labour supply decisions. As a result, we can distinguish between two cases. First, the individual's labour supply is completely inelastic and therefore given by the following expression

$$\ln y_{it} = (\Gamma_t + \psi_t' Z_{it} + \theta_t \alpha_i + u_{it}) + K_i \tag{2}$$

where K_i is the fixed number of hours of work of individual i per year, and it does not depend on his own wage. In this case, the features of earnings dynamics will be given by the following expression

$$Var(\ln y_{it}) = Var(\ln w_{it}) + Var(\ln K_i) \tag{3}$$

On the contrary, it could be that individuals show a significant labour supply behaviour because of the existence of significant responses in individuals' hours of work to shifts in their own wage rates. In this case, the cross-sectional variance of earnings inequality has the following form

$$Var(\ln y_{it}) = Var(\ln w_{it}) + Var(\ln h_{it}) + 2Cov(\ln w_{it}, \ln h_{it}) \tag{4}$$

Therefore, the responses of hours of work to wage changes (measured by the covariance in equation (4)) can increase or mitigate the effect of the dispersion in wages on the cross-sectional variance of earnings compared to the predictions obtained under the inelastic labour supply assumption. In this second case, we will have to consider an intertemporal individual labour supply model in order to capture the role of these responses of hours of work on the relationship between wage and earnings cross-sectional inequality.

As a first approximation to this question, we can compare predictions given by (3) and (4) with the information provided by Table 3 to determine which one better fits our data. In particular, we can observe that the cross-sectional variance of the log of earnings is smaller than the sum of the cross-sectional variance of the log of wages and the log of hours of work. Furthermore, the cross-sectional variance of the log of earnings is smaller than the cross-sectional of the log of wages. This suggests that there exists a negative cross-sectional covariance between the log of wages and the log of hours of work. To improve our understanding about this issue and identify the potential existence of a significant behaviour of the individual's hours of work, we need to incorporate a structural model of labour supply to our specification. However, first of all, we focus on the estimation of the wage equation and the decomposition of the wage cross-sectional inequality in permanent and transitory components.

3.2 Wage Model Estimation

Given the theoretical wage model specified in (1), we now specify the econometric model that we will fit to the empirical covariance structure of wages. In particular, we account for serial correlation in transitory wages and for measurement errors in observed wages. The empirical specification is given by the following expression,

$$\ln w_{it}^{*} = \ln w_{it} + \eta_{it} = \Gamma_t + \psi_t^{'} Z_{it} + \theta_t \alpha_i + u_{it} + \eta_{it} \tag{5}$$

where

$$u_{it} = \rho u_{it-1} + v_{it},$$

$\ln w_{it}^{*}$ represents the log of the observed wage, $\ln w_{it}$ is the log of the theoretical wage given by equation (1), and η_{it} captures the existence of measurement errors in wages and it is assumed to be white noise. The unobserved transitory component u_{it} follows a first-order autoregresive process where the error term v_{it} displays a zero mean and time-varying variance $\sigma_{v_t}^2$. This time-variation implies that transitory shocks may have different effects on wage inequality over time. For identification purposes the following assumptions are considered,

$$E(\alpha_i) = E(u_{it}) = E(\eta_{it}) = E(\alpha_i u_{it}) = E(\alpha_i \eta_{it}) = E(u_{it}\eta_{it}) = 0, \tag{6}$$

with the permanent scaled-factor for the first period θ_{94} normalized to one.

This specification allows us to quantify the contributions of observable and unobservable components of wages to changes in the overall cross-section wage inequality through a standard variance decomposition. In particular, we obtain the (OLS) predictions of the log of wages for each year from (5) using as regressors a set of observables characteristics, Z.[6] Then, we compute the variance of these predictions, and we get the following

[6] We include as regressors in the wage equation two dummies for the education level (secondary and graduate studies, respectively), the age and age squared, a dummy for the marital status, the specific experience and the specific experience squared, a set of dummies for the size and the economic activity of the firm, and the region. Estimation results from these regressions are available upon request.

decomposition

$$Var(\ln w_{it}^*) = Var(\widehat{\psi}_t^{'} Z_{it}) + Var(\xi_{it}^*) \tag{7}$$

where $\widehat{\psi}_t^{'}$ is the vector of estimated (OLS) returns to observable characteristics and ξ_{it}^* is the log wage OLS residual (which includes the effects of the unobservables). Notice that this decomposition of the overall cross-section wage inequality into between-group inequality (due to observables) and residual inequality (due to unobservables) relies on the orthogonality of the predicted values ($\widehat{\psi}_t^{'} Z_{it}$) and the residuals ($\xi_{it}^*$) in an OLS regression. Table 4 reports the results of this decomposition. As we can see, even though the between-inequality increases, the overall inequality decreases for the whole period due to the decrease in the within-inequality. Besides, both components represent on average 50 percent of the overall inequality during the whole period 1994-1998. This implies that our observable individual's characteristics explain about one half of the wage inequality each year and the returns of the unobserved skills, transitory shocks, measurement errors and estimation error count for the other one half. From now on we will focus on the analysis of the decomposition of wage residual inequality since this is the part of the inequality that can not be explained by the observables. Therefore, we consider the following wage equation

$$\xi_{it} = \Gamma_t + \theta_t \alpha_i + u_{it} + \eta_{it} \tag{8}$$

where ξ_{it} is the unobserved component of the log of wage, that is estimated by the log wage OLS residual, and u_{it} follows a first order autoregresive process, $u_{it} = \rho u_{it-1} + v_{it}$, where v_{it} displays a time-varying variance $\sigma_{v_t}^2$. The parameters of interest to be estimated are θ_t, the variance of the permanent component of wage (σ_α^2), the time-varying variance of the transitory shock ($\sigma_{\nu_t}^2$), the variance of the measurement error (σ_η^2) and the autoregresive coefficient in the transitory component (ρ).

After removing the effect of the aggregate terms, we obtain the following set of theoretical moments conditions implied by (6) and (8),

$$Cov(\xi_{it}, \xi_{it+k}) = \begin{matrix} \theta_t^2 \sigma_\alpha^2 + \sigma_{u_t}^2 + \sigma_\eta^2 & k = 0 \\ \theta_t \theta_{t+k} \sigma_\alpha^2 + \rho^k \sigma_{u_t}^2 & |k| > 0 \end{matrix} \quad (t = 1, ..., T) \tag{9}$$

where $\sigma_{u_t}^2 = \rho^2 \sigma_{u_{t-1}}^2 + \sigma_{v_t}^2$. Notice that for $t = 1$, $\sigma_{v_1}^2$ can not be identified. Instead of this parameter, we will estimate $\sigma_{u_1}^2$ that is equal to $\rho \sigma_{u_0}^2 + \sigma_{v_1}^2$.

Table 3. Cross-Sectional Variances, 1994-1998.

	Wage$^{(1)}$	Earnings$^{(2)}$	Hours$^{(3)}$
1994	0.2466	0.2236	0.03399
1995	0.2432	0.2270	0.0347
Percentage change (1994-1995)	(-1.4071)	(1.5050)	(1.9865)
1996	0.2635	0.2445	0.0328
Percentage change (1995-1996)	(8.3392)	(7.7058)	(-5.2842)
1997	0.2433	0.2137	0.0379
Percentage change (1996-1997)	(-7.6522)	(-12.6017)	(15.4338)
1998	0.2398	0.2128	0.03398
Percentage change (1997-1998)	(-1.4698)	(-0.4193)	(-10.3546)
Percentage change (1994-1998)	(-2.8088)	(-4.8509)	(-0.040)

Note: (1) Log of hourly wage.(2) Log of annual earnings. Wages and earnings have been deflacted by the annual mean of the Consumer Price Index (CPI) base 1992 and are expressed in euros.Number of observations, 1456 for 1994, 1413 for 1995, 1338 for 1996, 1676 for 1997 and 1665 for 1998.

Table 4. Wage Inequality Decomposition (1994-1998).$^{(1)}$

	1994	1995	1996	1997	1998
Overall Inequality	0.247	0.243	0.263	0.243	0.240
Between-Group Inequality	0.119	0.115	0.131	0.131	0.124
Percentage Explained (%)	48.202	47.496	49.800	53.763	51.665
Residual Inequality	0.128	0.128	0.132	0.112	0.116
Percentage Residual (%)	51.797	52.504	50.200	46.237	48.334
Sample Size	1456	1413	1338	1676	1665

Note:(1) Inequality is measured by the sample variance of the log hourly wage for each year. Results are based on OLS regressions of the log of wage on two dummies for the education level (secondary and graduate studies, respectively), the age and age squared, a dummy for the marital status, the specific experience and the specific experience squared, a set of dummies for the size and the economic activity of the firm, and the region.

This model is estimated using Generalized Method of Moments (GMM) techniques. As it is commonly known in the literature (Chamberlin (1984), Newey (1985), Abowd and Card (1989), Altonji and Segal (1996)), GMM estimators display generally desirable asymptotic properties. In addition to this, assumptions such that innovations are normally distributed are not necessary. Given our interest in the covariance structure of wages, we follow Abowd and Card (1989) and Dickens (2000) and minimize the weighted distance between the variance-covariance sample moments and the theoretical moments implied by the model in (8) and (6). Since we end up with 15 moment conditions and 12 parameters, the model is overidentified. In this case, different GMM estimators are obtained for different weighting matrices. In particular, we estimate by Equally Weighted Minimum Distance (EWMD) because it displays good properties in small samples applications (see Altonji and Segal (1996) for a further discussion).[7]

Table 5 summarizes the estimation results. First of all, we should point out that the value for the GMM goodness-of-fit statistic (6.625 with 3 degrees of freedom) indicates that this econometric specification represents a reasonable statistical description of our wage data. Furthermore, most the parameters are highly significant. The autocorrelation coefficient for the transitory wage shock presents a value by far less than one (0.554), which indicates that transitory shocks to wages do not show very high persistence and that their effect will decline rapidly over time. The permanent factor-loading θ_t shows little variation along years. Specifically, it slightly increases for 1995 and then decreases over the rest of the period. Based on these estimates, Table 6 reports the predictions for the variance of wage residuals and the relative contributions of permanent components and transitory components, and measurement errors to this variance. First of all, we compare the sample variance and the predicted variance that appear in the first two rows of Table 6. In particular, we can observe that our model underestimates slightly the cross-sectional variance from 1995 to 1998 and overestimates the decrease in the variance for the overall sample period (-9.3 for the actual variance and -15.32 for the predicted variance). Regarding the factor decomposition, we can see that transitory and permanent factors represent on average very similar contributions to the predicted variance (32.37

[7] In particular, Altonji and Segal (1996) show that eventhough Optimal Minimum Distance (OMD) estimator displays better asymptotic properties, it presents a downward bias in finite samples. Therefore, it is dominated by EWMD in small sample applications.

Table 5. Wage Model Estimates.

Parameter	EWMD	Parameter	EWMD
θ_{95}	1.089***	$\sigma^2_{v_2}$	0.016
	(0.149)		(0.013)
θ_{96}	1.002***	$\sigma^2_{v_3}$	0.032***
	(0.231)		(0.013)
θ_{97}	0.945***	$\sigma^2_{v_4}$	0.018*
	(0.207)		(0.013)
θ_{98}	0.938***	$\sigma^2_{v_5}$	0.022*
	(0.186)		(0.014)
σ^2_{α}	0.038***	σ^2_{η}	0.043***
	(0.016)		(0.011)
$\sigma^2_{u_1}$	0.047**	ρ	0.554**
	(0.012)		(0.308)
Goodness-of-fit	6.625	p-value	0.085

Note: Standard errors in parentheses. Degrees of freedom for goodness-of-fit statistics, 3. Number of observations, 2814 men with a total of 7548 individual-year observations. (*) Significant at 10%, (**) significant at 5%, (***) significant at 1%.

Table 6. Wage Inequality Predictions.

	EWMD Estimation.						
	1994	1995	1996	1997	1998	Average	94-98 change (%)
Sample Variance	0.1277	0.1276	0.1322	0.1124	0.1158	0.1231	-9.2989
Predicted Variance	0.1279	0.1187	0.1225	0.1080	0.1083	0.1171	-15.3197
Predicted Permanent Variance	0.0382	0.0453	0.0383	0.0341	0.0336	0.0379	-12.0410
Explained by Permanent Factors (%)[1]	29.8350	38.1834	31.2852	31.5443	30.9902	32.3676	3.8719
Predicted Transitory Variance	0.0467	0.0303	0.0412	0.0309	0.0317	0.0362	-32.0968
Explained by Transitory Factors (%)[1]	36.5373	25.5604	33.6022	28.6144	29.2984	30.7226	-19.8122
Predicted M. Errors Variance	0.0430	0.0430	0.0430	0.0430	0.0430	0.0430	0
Explained by Measurement Error (%)[1]	33.6277	36.2562	35.1126	39.8413	39.7113	36.9098	18.0913

Note: (1) Predicted relative weight of each component on the predicted residual wage inequality.

and 30.72 percent, respectively). However, their year-to-year fluctuations present different patterns. Whereas permanent components have slightly increased their relative relevance for the overall period, transitory factors have become substantially less important in explaining residual inequality. In particular, the relative weight of permanent component have increased about 4 percent, and the relative weight of transitory components have decreased about 20 percent. As a result, whereas in 1994 the relative contribution to the predicted variance is substantially larger for transitory factors, at the end of the sample period permanent factors are slightly more important. In addition to this, we should remark that the reduction in the predicted residual inequality is mostly explained by the substantial decrease in the transitory variance in the overall period. As Cervini and Ramos (2006) state, this reduction in earnings inequality due to transitory factors could be related to the labour market reforms implemented in 1994 and 1997 in Spain that pushed towards a decrease in temporary employment. With respect to the importance of measurement errors in wages, we have assumed that the variance of the term η_{it} is constant overtime. Our results show that this variance is very large and that a substantial part of the predicted residual variance is attributable to measurement error (about 37 percent on average). Furthermore, the relative importance of this component has increased considerably during the overall period.

3.3 A Life-Cycle Model of Individual Labour Supply

As we discussed at the end of Section 3, the decomposition of the variance of the log of earnings does not fit equation (3). In particular, we observed that the cross-sectional variance of the log of earnings is smaller than the sum of the cross-sectional variance of the log of wages and the log of hours of work suggesting the existence of a negative cross-sectional covariance between the log of wages and the log of hours of work. This suggests that individuals' responses to wage changes could play some "role" in the evolution of individual earnings inequality. In order to capture and quantify this "role", we use a theoretical intertemporal individual labour supply model following Hyslop (2001).[8] In particular, we specify an intertemporal model à la MaCurdy (1981). This structural model is based in

[8] Specifically, Hyslop (2001) proposes a structural intertemporal family labour supply model to analyse the relationship between household labour supply behaviour and family earnings inequality.

the following maintained assumptions. First, we assume that individual's time-preferences are represented by an utility function separable in labour supply and consumption. Moreover, preferences are intertemporal additively separable. Second, wages are assumed to be exogenous with respect individual's decisions and follow the process given by equation (1). Finally, we consider a deterministic framework (non-uncertainty) in the sense that individuals have perfect foresight about wages, nonlabour income and prices. As a result, each individual solves the following finite lifetime maximization problem:

$$\text{(P.1)} \qquad \underset{\{h_{it},c_{it}\}_{t=1}^{L}}{Max} \; v(h_i, c_i) = \sum_{t=1}^{L} \frac{1}{(1+\delta)^t} (\mu_{1it} c_{it}^{\beta_1} - \mu_{2it} (\frac{h_{it}}{\overline{H}})^{\beta_2})$$

$$s.t \qquad \sum_{t=1}^{L} \frac{1}{(1+r)^t} \left\{ w_{it} h_{it} + y_{it}^N - p_t c_{it} \right\} = 0$$

where the subscripts i denotes the individual, t denotes the period, and h_{it} and c_{it} represent the number of hours of work and consumption for individual i in period t, respectively. Regarding preferences components, we capture individual heterogeneity through μ_{1it} and μ_{2it}, that are age-specific modifiers of "tastes". Such variables depend on all of consumer i's characteristics which plausible affect his preferences at period t. These characteristics may include variables as the individual's education level, children and age. Besides, $\overline{H}$ represents the maximum number of hours of work, β_1 and β_2 are time-invariant parameters common across workers, with $0 < \beta_1 < 1$ and $\beta_2 > 1$, and δ is the discount factor or the time preference rate. The budget constraint defines incomes and expenditures at each period where w_{it} represents the individual's hourly true real wage in period t, that is assumed to follow the process specified by (1), and y_{it}^N, p_t and r are the nonlabour income, the price of goods, and the real interest rate, respectively.

Solving the first order conditions for individual's hours of work and assuming interior solutions,[9] we get the so-called $\lambda-$constant intertemporal labour supply for individual i in period t, that takes the following form

$$\ln h_{it} = \ln \overline{H} + \gamma \left[\ln \lambda_i + \ln w_{it} - t \ln \psi - \ln Y_{2it} - \ln \beta_2 \right] \tag{10}$$

where λ_i is the marginal utility of wealth for individual i (the Lagrange multiplier associated to the budget contraint), $\psi = \frac{1+\delta}{1+r}$ and $\gamma = \frac{1}{\beta_2 - 1}$. Notice that the wage term

[9] This assumption seems natural in this context since we are analysing the labour supply behaviour of working males.

that appears in (10) is the true wage and is given by (1). Therefore, the parameter γ is the intertemporal substitution elasticity and it is necessarily positive since it measures the response of hours of work to evolutionary changes in wages over the life cycle, holding constant the marginal utility of wealth, λ. In particular, it reflects how individuals allocate their working time toward those years where their wages are higher in their lifetime. We assume that $r = \delta$, so that, $\psi = 1$, and that "tastes" for work are randomly distributed over the population according to the equation $\ln \mu_{2it} = \varrho_t' X_{it} + \sigma_i - u_{it}^*$, where X_{it} is a vector of observable individual characteristics that affect hours of work, σ_i is a permanent component and u_{it}^* is a time-varying error term with zero mean.[10] Then, we get,

$$\ln h_{it} = \ln \overline{H} + \gamma \varrho_t' X_{it} + \gamma \left[\ln \lambda_i - \sigma_i - \ln \beta_2\right] + \gamma \ln w_{it} + \gamma u_{it}^* \tag{11}$$

Since individual's annual earnings are computed as the product of annual hours of work and the hourly wage, we obtain the following structural equation for individual's earnings:

$$\ln y_{it} = \ln w_{it} + \ln h_{it} = \ln \overline{H} + \gamma \varrho_t' X_{it} + \gamma F_i + (\gamma + 1) \ln w_{it} + \gamma u_{it}^* \tag{12}$$

where $F_i = (\ln \lambda_i - \sigma_i - \ln \beta_2)$ and represents a time invariant component unique to individual i that contains all the unobserved permanent components that affect labour supply decisions. Given (11) and (12), we should point out the most salient features of this model. First of all, given that the parameter γ measures the individual's intertemporal labour supply elasticity, there will exist a significant labour supply behaviour for individual i as long as this parameter is nonzero. If this is the case, individual's labour supply behaviour will be determined by his own wage life-time stream in two ways: first, the number of hours of work will depend for each period on its own contemporaneous wage (substitution effect); second, it will also depend on F_i, that summarizes through $\ln \lambda_i$ all of the retrospective and prospective information about real wages and non-labour income relevant to the individual i's choices. However, if γ is zero, then each individual will work a fixed number of hours, given by $\ln \overline{H}$, and the earnings equation will fit the "no labour supply" earnings process given in equation (2). Secondly, we consider a deterministic framework, that implies that individuals have perfect foresight with respect

[10] This decomposition is proposed by MaCurdy (1981).

to their vectors of lifetime wages, prices and nonlabour income. Therefore, the model does not reflect individual responses to unanticipated wage variations. On the contrary, agents fully anticipate all wage changes over the period, and therefore, any labour supply response represents a purely intertemporal substitution effect along the $\lambda-$constant intertemporal labour supply function.

4 Wages and Earnings Model Estimation

From equation (1) and (12), we derive the following empirical specification for the joint model of wages and earnings:

$$\ln w_{it}^{*} = \ln w_{it} + \eta_{it} = \Gamma_t + \psi_t^{'} Z_{it} + \theta_t \alpha_i + u_{it} + \eta_{it} \tag{13a}$$

$$\ln y_{it}^{*} = \ln y_{it} + \varepsilon_{it} = \ln \overline{H} + \varphi_t^{'} Y_{it} + \gamma F_i + (\gamma + 1)(\Gamma_t + \theta_t \alpha_i + u_{it}) + \gamma u_{it}^{*} + \varepsilon_{it} \tag{13b}$$

where Y_{it} is a vector of individual observable characteristics that affect earnings and includes Z_{it} and X_{it}. The term ε_{it} captures measurement errors in individual earnings and is assumed to be white noise. For identification purposes, we assume in addition to (6) that

$$E(\varepsilon_{it}\alpha_i) = E(\varepsilon_{it}u_{it}) = E(\varepsilon_{it}F_i) = E(\varepsilon_{it}u_{it}^{*}) = E(F_i u_{it}^{*}) = E(u_{it}u_{it}^{*}) = E(\alpha_i u_{it}^{*}) = 0, \tag{14}$$

with the permanent scaled-factor for the first period θ_{94} normalized to one. Therefore, we only allow for correlation between the permanent components in wages and hours of work (α_i and F_i, respectively) and between the measurement errors in wages and earnings.[11] Notice that γu_{it}^{*} (the transitory component of the "tastes" of work) and ε_{it} can not be separately identified. Therefore, we will assume a particular stochastic process for the whole component $\gamma u_{it}^{*} + \varepsilon_{it}$, that will capture the transitory determinants of hours of work and measurement errors.

[11] Remember that hourly wages have been computed as the ratio between annual earnings and annual hours of work. Therefore, η_{it} and ε_{it} could be correlated between each other.

Using equation (13b), we can quantify the contributions of observable and unobservable components of earnings to changes in the overall cross-section inequality through the same variance decomposition that we performed for wages. In particular, if we obtain the (OLS) predictions of the log of earnings for each year using as regressors the set of observables characteristics, Y,[12] and we compute the variance of these predictions, we get the following expression

$$Var(\ln y_{it}^{*}) = Var(\widehat{\varphi}_t^{'} Y_{it}) + Var(\varsigma_{it}^{*}) \tag{15}$$

where $\widehat{\varphi}^{'}$ is the vector of estimated (OLS) returns to observable characteristics and ς_{it}^{*} is the log earnings OLS residuals for each individual i in each period t (which includes all the effects of the unobservables). Table 7 reports the results of this decomposition. Given these results, we can observe that both components of the overall inequality in log earnings have decreased in the period although such pattern do not hold for all years. Besides, the decrease in residual inequality is larger than in between-group inequality. Regarding the relative importance of each component, we should notice that observable variables count on average for sligthly more than 50 percent, and they have increased their contribution in the period from 50 percent to 52 percent. However, similar to the case of wage inequality, both components play rather the same role in explaining the overall earnings inequality. As we pointed out below, we now focus on the analysis of the decomposition of earnings residual inequality since this is the part of the inequality that can not be explained by our observables.

Therefore, we consider the following system for the unobserved components of both the log of wage and the log of earnings,

$$\xi_{it} = \Gamma_t + \theta_t \alpha_i + u_{it} + \eta_{it} = \ln w_{it}^{+} + \eta_{it} \tag{16a}$$

$$\varsigma_{it} = \ln \overline{H} + \gamma F_i + (\gamma + 1) \ln w_{it}^{+} + \zeta_{it} \tag{16b}$$

where ξ_{it} and ς_{it} are the unobserved component of the log of wages and the log of earnings, respectively, that are estimated by the corresponding OLS residuals, $\ln w_{it}^{+}$ is

[12] We consider as regressors in the earnings equation for each year the same variables that we included in the wage equation, and two additional variables, the number of children less that 6 year old and the number of children between 6 and 18. Estimation results from these regressions are available upon request.

the unobservable theoretical wage that is equal to $\Gamma_t + \theta_t\alpha_i + u_{it}$, and ζ_{it} is equal to $\gamma u^*_{it} + \varepsilon_{it}$ and assumed to be white noise. The parameters of interest to be estimated are θ_t, the variance of the permanent component of wage (σ^2_α), the time-varying variance of the transitory shock ($\sigma^2_{\nu_t}$), the variance of ζ_{it} (σ^2_ζ), the variance of the measurement error in wages (σ^2_η), the covariance between the permanent components of wages and earnings ($\sigma_{\alpha,F}$), the individual' labour supply elasticity (γ), the autoregresive coefficient in the wage transitory shocks (ρ), and the covariance between the measurement error term in wages η_{it} and ζ_{it} ($\sigma_{\eta,\zeta}$). Again, notice that for $t = 1$, $\sigma^2_{v_1}$ can not be identified. Instead of this parameter, we will estimate $\sigma^2_{u_1}$ that is equal to $\rho\sigma^2_{u_0} + \sigma^2_{v_1}$. For identification purposes, we normalize σ^2_F=1.[13]

From the system (16a) and (16b), we obtain the set of theoretical moment conditions for wages given by (9), and the following set of theoretical moments conditions for earnings derived from (16b)

$$Cov(\varsigma_{it}, \varsigma_{it+k}) = \begin{cases} \gamma^2\sigma^2_F + (1+\gamma)^2 Var(\ln w^+_{it}) + Var(\zeta_{it}) + 2\gamma(1+\gamma)\theta_t\sigma_{\alpha,F} & k=0 \\ \gamma^2\sigma^2_F + (1+\gamma)^2 cov(\ln w^+_{it}, \ln w^+_{it+k}) + \gamma(1+\gamma)(\theta_t + \theta_{t+k})\sigma_{\alpha,F} & |k|>0 \end{cases} \quad \forall t \tag{17}$$

where the theoretical moments predictions for the true unobserved wages, $\ln w^+_{it}$, are given by (9) excluding the measurement error component. In addition to this, we use the cross-moment conditions of wages and earnings that are given by

$$Cov(\xi_{it}, \varsigma_{it+k}) = \begin{cases} \gamma\theta_t\sigma_{\alpha,F} + (1+\gamma)^2 Var(\ln w^+_{it}) + \sigma_{\eta,\zeta} & k=0 \\ \gamma\theta_t\sigma_{\alpha,F} + (1+\gamma)^2 Cov(\ln w^+_{it}, \ln w^+_{it+k}) & |k|>0 \end{cases} \quad \forall t \tag{18}$$

Similarly to the wage equation, we estimate the present model by EWMD. Specifically, we fit the 55 second moment predictions implied by our econometric model (16a) and (16b) to the empirical covariance structure of these variables for our sample of continuously working males.

[13] Results from the estimation of parameters σ^2_F, $\sigma_{\alpha F}$, and γ based on simulations of the joint model of wages and earnings show that these parameters are very imprecisely estimated. In particular, the standard deviation of the estimate of σ^2_F for different simulated samples turned out to be very large unless sample sizes were considerably large. This points to an identification problem of the parameters of the model. Therefore, we decided to set σ^2_F equal to 1.

Table 7. Earnings Inequality Decomposition (1994-1998).(1)

	1994	1995	1996	1997	1998
Overall Inequality	0.224	0.227	0.244	0.214	0.213
Between-Group Inequality	0.112	0.109	0.125	0.118	0.111
Percentage Explained (%)	50.074	48.240	51.224	55.324	51.962
Residual Inequality	0.112	0.117	0.119	0.095	0.102
Percentage Residual (%)	49.926	51.760	48.776	44.676	48.038
Sample Size	1456	1413	1338	1676	1665

Note:(1) Earnings inequality is measured by the sample variance of log earnings for each year. Results are based on OLS regressions of the log of earnings on two dummies for the education level (secondary and graduate studies, respectively), the age and age squared, a dummy for the marital status, the specific experience and the specific experience squared, a set of dummies for the size and the economic activity of the firm, the region, and two fertility variables (children less than 6, and children between 6 and 18).

Table 8. Wage and Labour Supply Model Estimates.

θ_{95}	1.058***	$\sigma^2_{v_4}$	0.014
	(0.095)		(0.012)
θ_{96}	0.999***	$\sigma^2_{\nu_5}$	0.019*
	(0.151)		(0.012)
θ_{97}	0.958***	σ^2_{η}	0.045***
	(0.122)		(0.010)
θ_{98}	0.971***	σ^2_{ζ}	0.021**
	(0.103)		(0.012)
σ^2_{α}	0.051***	$\sigma_{\alpha F}$	-0.089***
	(0.007)		(0.013)
$\sigma^2_{u_1}$	0.030***	ρ	0.459*
	(0.012)		(0.288)
$\sigma^2_{v_2}$	0.019**	γ	0.121***
	(0.009)		(0.006)
$\sigma^2_{v_3}$	0.029***	$\sigma_{\eta\zeta}$	0.024**
	(0.012)		(0.011)
$Goodness-of-fit$	55.455	$p-value$	0.042

Note: Standard errors in parentheses. Degrees of freedom for goodness-of-fit statistics, 39. Number of observations, 2814 men with a total of 7548 individual-year observations. (*) Significant at 10%, (**) significant at 5%, (***) significant at 1%.

Table 8 presents the estimation results for the joint model of unobserved wages and earnings. Most of the parameter estimates for the wage equation are highly significant and pretty similar to those in Table 5, except that the autocorrelation parameter in the transitory component of wages is smaller in this case. With respect to the parameter estimates from the earnings equation, we can see that the covariance between the permanent wage and earnings components ($\sigma_{\alpha F}$) is negative and very significant. Given the estimate for σ_{α}^2 and the fact that we have normalized σ_F^2 to 1, this implies a correlation between both components equal to -0.394. Similarly, we can infer from the estimation results that the correlation between the wage measurement error term η_{it} and the term ζ_{it} in the earnings equation is substantially high (0.781).[14] Finally, with respect to the labour supply behaviour, we can see that our parameter γ is significantly positive but small (0.121). This result is in line with the evidence shown by MaCurdy (1981) and Altonji (1986) that arrive at estimates of this intertemporal substitution elasticity in the neighborhood of 0.10 and 0.40 using data drawn from PSID.[15] Therefore, this result that the hypothesis of "no labour supply" is clearly rejected by our data.

The GMM goodness-of-fit statistic (55.455, with 39 degrees of freedom) reflects that this wage and labour supply model does not provide a very good statistical fit to the data. However, as Hyslop (2001) notes this fit is comparable to other models fit to covariance structures in the literature (Dickens (2000), Hyslop (2001), Blundell et al. (2003)) and could be due to the fact that the labour supply model is extremely parsimonious.

5 Earnings Inequality Predictions

In this section, we use the estimation results obtained for the joint model of individuals' unobserved wages and earnings to compute the predicted cross-sectional variance of the log of earnings residuals along our sample period. The comparison of these predictions with the sample cross-sectional variance gives us information about the ability of

[14] Remember that the term ζ_{it} is equal to $\gamma u_{it}^* + \varepsilon_{it}$, where ε_{it} represents the measurement errors in earnings. Therefore, given that wages are computed by the ratio between annual earnings and annual hours of work, we should expect a positive correlation between η_{it} and ζ_{it}.

[15] In addition to this evidence, there are other papers in the literature that have obtained very imprecise estimates of the intertemporal labour supply elasticity (i.e. Abowd and Card (1989) for male labour supply and Arellano et al.(1999) for female labour supply).

our model to reproduce the empirical evidence shown by our sample. Furthermore, we quantify the relative contributions of permanent components, transitory components and the individuals' labour supply decisions on the levels and changes in residual earnings inequality.

Table 9 reports the results of this quantitative analysis. Specifically, the first two rows report the level and the change in the sample variance and the predicted variance of the log of earnings residuals between 1994 and 1998, respectively. We can observe that these two elements turn out to be pretty similar for all years. This shows that the model for wages and earnings given by (16a) and (16b) provides a reasonable approximation to the empirical evidence in our sample. In particular, we can see that the data and the predictions obtained from the model show a reduction of 8.5-10.5 percent in the residual earnings inequality for the overall period eventhough there are some year-to-year fluctuations. In addition to this, this decrease is smaller than the reduction experienced by the residual wage inequality shown in Table 6. Rows 3-6 present the decomposition of the predicted residual earnings inequality into components attributable to permanent factors and transitory factors. In particular, results in rows 3 and 5 show that permanent inequality and transitory inequality[16] decrease over the period. Moreover, this reduction is substantially larger for the part of the cross-sectional variance that is explained by transitory factors. With respect to the relative contribution of these two types of factors, we should notice from rows 4 and 6 that both of them play a very similar role in the levels of earnings inequality (around 50 percent on average). However, permanent factors have increased slightly their relative contribution over the period by almost 6 percent. This change in the factor composition of earnings inequality together with the significant decrease in the transitory inequality is in line with the evidence shown by Cervini and Ramos (2006) for a sample of male full-time employees from the ECHP. In particular, they find that in 1994 the transitory component played a somewhat larger role but, over the next years, earnings dispersion became more persistent and less transitory. As they argue, this could show the effectiveness of the labour market reforms undertaken in 1994

[16] Permanent inequality is determined by the terms involving wage permanent components and F_i. Transitory inequality is determined by the wage transitory components and ζ_{it}. As we mentioned above, ζ_{it} is equal to $\gamma u^*_{it} + \varepsilon_{it}$ and u^*_{it} and ε_{it} can not be separately identified. Therefore, we consider the measurement error term ε_{it} as part of the transitory components.

and 1997 in Spain which aimed for a reduction of temporary employment. In fact, if we decompose the residual earnings inequality change over the period into permanent factors and transitory factors, our results suggest that a quarter of this variation is explained by the reduction in the permanent inequality and three quarters is due to transitory factors.[17]

Finally, we now refer to the relative contribution of individuals' labour supply responses to the cross-sectional residual earnings inequality. This analysis requires the computation of a counterfactual for the residual earnings inequality under the assumption of non-existence of a significant labour supply behaviour. In particular, we compute the predicted residual earnings inequality assuming that the intertemporal substitution elasticity γ is zero. If we apply this condition to equation (16b), we obtain the following

$$\varsigma_{it} = \ln \overline{H} + \ln w_{it}^{+} + \varepsilon_{it} \tag{19}$$

Notice from this equation that we need to estimate σ_{ε}^2 to compute the predicted variance of earnings residuals under the counterfactual. However, as we mentioned above, σ_{ε}^2 and $\sigma_{u^*}^2$ can not be identified separately. Therefore, the exact value for the predicted earnings inequality given by (19) can not be computed from our estimates and only bounds can be provided. Specifically, the lower bound is given by the predictions obtained assuming that there is not measurement error in earnings. Then, in this case, ζ_{it} is equal to γu_{it}^* and the lower bound is given by the predicted cross-sectional variance of $\ln w_{it}^{+}$. The upper bound is contructed assuming that ζ_{it} is given exclusively by the measurement error in earnings. Therefore, ζ_{it} is equal to ε_{it} and the upper bound it is determined by the predicted cross-sectional variance of $\ln w_{it}^{+} + \zeta_{it}$. Rows 7 and 8 in Table 9 present the results for these upper and lower bounds of the predicted residual earnings inequality in the absence of a significant labour supply behaviour, respectively. In addition to this, rows 9 and 10 show the lower and upper bound for the contribution of individuals' labour supply behaviour to residual earnings inequality, respectively.[18] In particular, we can see that

[17] In particular, results in Table 9 show that the predicted residual earnings inequality has decreased over the period by 10.5 percent. If we decompose this reduction into permanent and transitory components, we obtain that the decrease due to permanent factors and transitory factors is 2.5794 and 7.9137, respectively. Therefore, our model predicts that a quarter of the reduction in residual earnings inequality (10.5 percent) is explained by the reduction in the permanent inequality and three quarters is due to transitory factors.

[18] The relative contribution of individuals' labour supply has been computed as the relative weight on the predicted residual earnings inequality of the difference between the residual earnings inequality

Table 9. Earnings Inequality Predictions

EWMD Estimation.

	1994	1995	1996	1997	1998	Average	94-98 change (%)
Sample Variance	0.1116	0.1174	0.1192	0.0954	0.1022	0.1091	-8.4416
Predicted Variance	0.1137	0.1138	0.1187	0.0978	0.1018	0.1091	-10.4931
Predicted Permanent Variance	0.0548	0.0610	0.0547	0.0505	0.0519	0.0546	-5.3517
Explained by Permanent Factors (%)$^{(1)}$	48.1973	53.6343	46.0972	51.6534	50.9658	50.1096	5.7441
Predicted Transitory Variance	0.0589	0.0527	0.0640	0.0473	0.0499	0.0545	-15.2766
Explained by Transitory Factors (%)$^{(1)}$	51.8027	46.3657	53.9028	48.3466	49.0342	49.8904	-5.3443
Predicted Variance under $\gamma = 0^{(2)}$	0.1027	0.1034	0.1062	0.0888	0.0922	0.0986	-9.8208
Predicted Variance under $\gamma = 0^{(3)}$	0.0813	0.0825	0.0853	0.0679	0.0713	0.0777	-12.3481
Labour Supply (%)$^{(2)}$	10.0554	9.0957	10.5040	9.1967	9.3798	9.6463	-6.7188
Labour Supply (%)$^{(3)}$	28.4649	27.4904	28.1366	30.6016	29.9475	28.9282	5.2084

Note: (1) Predicted relative weight of each component on the predicted residual earnings inequality. (2) Predicted cross-sectional variance of earnings and contribution of labour supply behaviour to the variance of earnings assuming that ζ_{it} is equal to ε_{it}. (3) Predicted cross-sectional variance of earnings and contribution of labour supply behaviour to the variance of earnings assuming that ζ_{it} is equal to γu^*_{it}.

Table 10. Predicted Cross-Sectional Variances, 1994-1998.

	Unobserved Theoretical Wage	Unobserved Theoretical Earnings$^{(1)}$
1994	0.0813	0.0928-0.1137
1995	0.0825	0.0928-0.1138
1996	0.0853	0.0978-0.1187
1997	0.0679	0.0768-0.0978
1998	0.0713	0.0808-0.1018

Note: (1) Lower and upper bounds.

the relative contribution of the individuals' labour supply range on average for the overall period between 9 and 28 percent. This means that in the absence of a significant labour supply behaviour, residual earnings inequality would have been lower in 9-28 percent. Therefore, this evidence suggests that intertemporal substitution effects in which each individual's hours of work react positively to variations in the own wage reinforce the transmission of wage inequality to earnings inequality. Related to this, we compare in Table 10 the predicted cross-sectional variance of the unobservable theoretical wages and the cross-sectional variance of the unobservable theoretical earnings. From the predictions of equation (4) and the existence of a significant and positive labour supply behaviour, we expect that the cross-sectional variance of the unobservable theoretical wages is smaller than the cross-sectional variance of the unobservable theoretical earnings. To confirm this, we should compute the predicted cross-sectional variance of $\ln w_{it}^{+} = \Gamma_t + \theta_t \alpha_i + u_{it}$ and the predicted cross-sectional variance $\varsigma_{it}^{+} = \ln \overline{H} + \gamma F_i + (\gamma+1) \ln w_{it}^{+} + \gamma u_{it}^{*}$. However, as we explained above, $\zeta_{it} = \gamma u_{it}^{*} + \varepsilon_{it}$ and u_{it}^{*} can not be separately identified from ε_{it} in our model. Therefore, we compute a lower and upper bound for this cross-sectional variance. The lower bound has been computed assuming that ζ_{it} is given exclusively by the measurement error in earnings. Therefore, ζ_{it} is equal to ε_{it} and the lower bound for the cross-sectional variance of the unobservable theoretical earnings is determined by the cross-sectional variance of $\gamma F_i + (\gamma + 1) \ln w_{it}^{+}$. On the other hand, the upper bound has been computed assuming that ζ_{it} is given exclusively by γu_{it}^{*}. Therefore, the upper bound is determined by the cross-sectional variance of $\gamma F_i + (\gamma+1) \ln w_{it}^{+} + \gamma u_{it}^{*}$. Results in Table 10 show that the cross-sectional variance of the unobserved theoretical wages is always smaller than the cross-sectional variance of the unobserved theoretical earnings. This confirms again that positive responses to wage rates changes enlarge earnings inequality with respect to wage inequality. This differs from the results in terms of the observable variables shown in Table 3 and section 3.1 where we saw that the cross-sectional variance of the log of earnings was smaller than the cross-sectional variance of the log of wages. Notice that this result could be driven by the importance of the variance of wages explained by our observables and the large relative contribution of the measurement error in wages (around 37 percent on average for the overall period).

Finally, we should remark that the previous results on earnings inequality predictions

predicted by the model and the residual earnings inequality predicted assuming $\gamma = 0$.

have been computed under the assumption that σ_F^2 is equal to one. As a robustness check of our results to this normalization, we estimate the joint model of wages and earnings assuming two extreme values for this parameter. Table 11 and Table 12 in the Appendix show the predictions and the decomposition of residual earnings inequality obtained when σ_F^2 is set to 0.01 and 5000, respectively. From these results, we conclude that this normalization assumption is not an important concern since predictions are very similar under these two alternative values.[19]

6 Conclusion

In this paper we document the evolution of wage and earnings inequality for a sample of Spanish continuously working males from the ECHP for the period from 1994 to 1998. In particular, we analyse the link between the dispersion in the wage rates and individual earnings inequality taking into account two labour market features that could affect this transmission. First, the distinction between permanent components and transitory components and their relative contribution to wage and earnings inequality. Second, the existence of significant responses in individuals' hours of work to shifts in their own wage rates. For this analysis we use an error component model for wages and earnings that is based on an intertemporal individual labour supply model à la MaCurdy (1981). This model is estimated by EWMD.

The evidence presented in this paper shows that cross-sectional inequality in wages and individual earnings have decreased during the sample period by almost 3 percent and 5 percent, respectively. When we decompose inequality in "between-group inequality" and "residual inequality", we obtain that our observables explain on average around 50 percent of the cross-sectional variance in wages and in individual earnings. Given the magnitude of the cross-sectional dispersion that can not be explained by our observable characteristics of individuals and their jobs, we focus on the variance-decomposition analysis of the residual inequality of wages and earnings. Our results suggest that both wage and

[19] The estimation of the model assuming different values for σ_F^2 shows that $\widehat{\sigma_{\alpha F}}$ gets larger in absolute value and $\widehat{\gamma}$ gets smaller as σ_F^2 increases. However, this does not affect the results for the predictions and the decomposition of inequality. Results for the parameter estimates of the model are available upon request.

earnings residual inequality have decreased in the overall period although the reduction for earnings has been smaller. Besides, the decomposition of the earnings inequality shows that both permanent variance and transitory variance have decreased between 1994 and 1998 but permanent factors have increased their relative contribution in explaining earnings inequality. This reflects that earnings dispersion have become more persistent and less transitory. Moreover, around three quarters of the reduction in earnings inequality is attributed to transitory factors. With respect to individual labour supply behaviour, we find a significant and positive intertemporal substitution elasticity of 0.121. This implies that residual earnings inequality would have been lower in 9-28 percent in the absence of a significant labour supply behaviour for males in our sample. Given that the predicted earnings variance has decreased around 10.5 percent in the overall period, this result confirms that male labour supply behaviour plays a substantial role in the evolution of earnings inequality.

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APPENDIX

Table 11 Earnings Inequality Predictions ($\sigma_F^2 = 0.01$)

EWMD Estimation.

	1994	1995	1996	1997	1998	Average	94-98 change (%)
Sample Variance	0.1116	0.1174	0.1192	0.0954	0.1022	0.1091	-8.4416
Predicted Variance	0.1134	0.1140	0.1181	0.0983	0.1019	0.1091	-10.0778
Predicted Permanent Variance	0.0523	0.0589	0.0521	0.0471	0.0488	0.0518	-6.8121
Explained by Permanent Factors (%)$^{(1)}$	46.1763	51.6456	44.1686	47.8556	47.8533	47.5399	3.6317
Predicted Transitory Variance	0.0610	0.0551	0.0659	0.0513	0.0532	0.0573	-12.8796
Explained by Transitory Factors (%)$^{(1)}$	53.8237	48.3544	55.8314	52.1444	52.1467	52.4601	-3.1157
Predicted Variance under $\gamma = 0^{(2)}$	0.1056	0.1061	0.1099	0.0917	0.0951	0.1017	-9.9900
Predicted Variance under $\gamma = 0^{(3)}$	0.0795	0.0799	0.0838	0.0656	0.0689	0.0755	-13.2809
Labour Supply (%)$^{(2)}$	6.8320	6.9125	6.8439	6.6984	6.7410	6.8056	-1.3315
Labour Supply (%)$^{(3)}$	29.9179	29.8740	29.0131	33.3156	32.4142	30.9069	8.3439

Note: (1) Predicted relative weight of each component on the predicted residual earnings inequality. (2) Predicted cross-sectional variance of earnings and contribution of labour supply behaviour to the variance of earnings assuming that ζ_{it} is equal to ε_{it}. (3) Predicted cross-sectional variance of earnings and contribution of labour supply behaviour to the variance of earnings assuming that ζ_{it} is equal to γu_{it}^*.

Table 12. Earnings Inequality Predictions ($\sigma_F^2 = 5000$)

EWMD Estimation.

	1994	1995	1996	1997	1998	Average	94-98 change (%)
Sample Variance	0.1116	0.1174	0.1192	0.0954	0.1022	0.1091	-8.4416
Predicted Variance	0.1133	0.1135	0.1177	0.0989	0.1023	0.1091	-9.6778
Predicted Permanent Variance	0.0557	0.0617	0.0555	0.0512	0.0526	0.0553	-5.6224
Explained by Permanent Factors (%)$^{(1)}$	49.1764	54.3857	47.1514	51.7545	51.3844	50.7705	4.4899
Predicted Transitory Variance	0.0576	0.0518	0.0622	0.0477	0.0497	0.0538	-13.6018
Explained by Transitory Factors (%)$^{(1)}$	50.8236	45.6143	52.8486	48.2455	48.6156	49.2295	-4.3444
Predicted Variance under $\gamma = 0^{(2)}$	0.1076	0.108299	0.1120	0.0929	0.0964	0.1034	-10.4008
Predicted Variance under $\gamma = 0^{(3)}$	0.0840	0.0847	0.0884	0.0693	0.0728	0.0799	-13.3166
Labour Supply (%)$^{(2)}$	5.0182	4.5868	4.8564	6.0894	5.7786	5.2659	15.1508
Labour Supply (%)$^{(3)}$	25.8152	25.3452	24.8720	29.9124	28.8038	26.9497	11.5771

Note: (1) Predicted relative weight of each component on the predicted residual earnings inequality. (2) Predicted cross-sectional variance of earnings and contribution of labour supply behaviour to the variance of earnings assuming that ζ_{it} is equal to ε_{it}. (3) Predicted cross-sectional variance of earnings and contribution of labour supply behaviour to the variance of earnings assuming that ζ_{it} is equal to γu_{it}^*.