WAGE ADJUSTMENTS, WORKER FLOWS, AND UNEMPLOYMENT

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Tese de Doutoramento em Economia

Orientada por
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Faculdade de Economia
Universidade do Porto
Maio, 2003

Acknowledgments

So many reasons to be grateful, so many people to be grateful to. First, and above all, thank You God for being who You are. Awesome! I am indebted for each person You brought into my life.

I am grateful to Professor Pedro Portugal for his enthusiasm, advice and encouragement in difficult times. I deeply appreciate his guidance during this research project and I deeply thank him for being present whenever I needed his help.

I am also grateful to Professor Mendes de Oliveira for his encouragement and his accurate comments.

A special thanks goes to Varejão and Helder for dozens of small 'nothings' that meant a lot to me. Varejão also commented some parts of this dissertation.

I have also to acknowledge the encouragement of my colleagues at CETE and, in particular, its director Margarida Ruivo. I am deeply grateful to Margarida for her willingness to help me, especially during the semester I taught Microeconomics with her. I also thank Pilar and Luís Delfim for writing the abstract in French.

My gratitude also goes to the Departamento de Estatística do Ministério da Segurança Social e do Trabalho and to its director, Dra. Maria João Rebelo, for allowing me to use the data from the *Social Audit*. I am also grateful to Manuel Arellano for kindly supplying me with the DPD program.

A word of gratitude is also due to the Faculdade de Economia do Porto for their institutional support. I, especially, would like to thank the current Director, Professor José da Silva Costa, and the President of the Scientific Council, Professor António Almodovar. A word of appreciation also goes to the staff of the library and of the computing department of the Faculdade de Economia do Porto.

This research was funded by Fundo Social Europeu e Estado Português under grant Concurso 2/5.3/PRODEP/2000, which I gratefully acknowledge. Financial support from Fundação para a Ciência e Tecnologia under research grant POCTI/ECO/35147/99 is also warmly acknowledged.

Finally, my last thoughts of gratitude go to my family and friends. In particular, I owe a very special acknowledgment to my parents for their unconditional support and to João and João Francisco for their encouragement and patience.

Abstract

This dissertation comprises four essays on wage adjustments. The first investigates real wage cyclicality using Portuguese data, addressing the issue of heterogeneity in wage responses to aggregate labor market conditions for workers' accessions and separations. The results reveal that entry wages are much more procyclical than current wages, which confirms previous empirical research consistent with the hypothesis of implicit shielding agreements between employers and workers in wage determination. Moreover, the empirical evidence found suggests that even micro-data estimates of real wage cyclicality may conceal a strong procyclical wage behavior, when heterogeneity on wage responses to aggregate conditions between employed workers and hirings and separations are not taken into account.

The second essay investigates wage determination at the firm level using a longitudinal panel of large firms. Adopting a rent-sharing model of wage determination two main objectives drive the investigation. The first is to evaluate the role and weight of insider and outsider forces in the process of wage determination. The second is to test empirically to what extent the existence of labor turnover costs creates market power that is exploited by incumbent workers through wage claims. The results indicate that wages seem to be very sensitive to firm-specific conditions and that insider-outsider considerations have significant empirical relevance for Portugal. In particular, the results reveal that firms where insider workers seem to have more market power tend to pay higher wages.

The third essay examines the impact of displacement on long-term earnings losses of displaced workers in Portugal, using a nationally representative longitudinal linked employer-employee data set. The results reveal that the effects of displacement on earnings are considerable in magnitude and quite persistent. The post-displacement earnings losses are mainly associated with the loss of tenure within the firm and, to a lesser extent, to the loss of sector-specific features. Furthermore, workers who experienced a spell of nonemployment are the most affected by job displacement. Finally, this study points to the importance of controlling for employers' characteristics in this type of wages-dynamic analysis, since there are systematic differences in earnings between displaced and non-displaced workers that stem from differences in firm characteristics. Ignoring them may confound the evaluation of the earnings losses.

In the fourth essay a simultaneous-equations model of firm closing and wages is developed in order to analyze how wages adjust to unfavorable shocks that raise the risk of displacement through firm closing, and to what extent a wage change affects the exit likelihood. The results show that the fear of job loss generates bargaining concessions instead of compensating differentials. A truly novel result that emerges from this study is that firms with a higher incidence of minimum wage earners are more vulnerable to adverse demand shocks due to their inability to adjust wages downward. In fact, minimum wage restrictions were seen to increase the failure rates.

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Chapter 1

Workers' Flows and Real Wage Cyclicality

1 Introduction

The cyclical behavior of real wages has been the subject of a large number of studies since the debate of Keynes (1939), Dunlop (1938), and Tarshis (1939). However, despite considerable research on this issue, the same vexed questions remain to be solved. Are real wages countercyclical or procyclical? Flexible price-aggregate demand models predict that real wages should move countercyclically as employment adjusts along a negatively sloped aggregate demand for labor schedule. Real business cycle models, on the other hand, predict procyclical movements as workers substitute work for leisure along dynamic labor supply schedules.

Brandolini (1995) presents an exhaustive literature review of the studies on real wage cyclicality published from the 1920s onwards. Abraham and Haltiwanger (1995) also present a review of the empirical evidence on real wage cyclicality, stressing the issue of the sensitivity of the results to the measurement of nominal wages, nominal prices, and cyclical conditions, as well as to frequency, time period, and econometric specification.

The studies based on aggregate data show some ambiguous results. In this case, the best conclusion is that the choice of the time period analysis, price deflator, and cyclical indicator, as well as the choice between wage rates and average earnings (including overtime or not), may substantially affect the estimates of real wage cyclicality [Abraham and Haltiwanger (1995)].

One reason why these studies have reached no definitive conclusions resides in the fact that they have been performed at the aggregate level. In particular, they have ignored the changes in the composition of the workforce over the cycle. The presence of compositional effects has attracted much attention in the last years and recent micro-data studies based on panel data for the U.S. showed that composition bias plays an important role on real wage behavior along the business cycle [see, for example, Mitchell *et al.* (1985), Bils (1985), Keane *et al.* (1988) and Solon *et al.* (1994)]. In fact, cyclical changes in the composition of the work force may induce a countercyclical bias in the aggregate real wage. Aggregate measures of real wages tend to give more weight to low-skill workers during expansions than during recessions. The argument is that if less-skilled workers are more vulnerable to layoff, they will account for a smaller share of employment in recessions than in expansions.

As stated in Blundell *et al.* (2000), "Aggregate figures for real wage growth are used extensively in policy debate to analyze changes in the well-being of workers over time and to compare different groups of people both within and

across countries. However, if participation (employment) rates change across time periods or across the groups used in these comparisons, then aggregate real wages may give a misleading impression of changes in the structure of real wages facing individual workers." According to them, aggregate real wages are shown to contain three important bias terms: one that describes the dispersion of wages, the second reflecting the distribution of working hours, and the third captures the effects of composition changes in the selected sample of workers. Using data from the U.K. Family Expenditure Survey (FES) from 1978 to 1996, they have shown that those three terms play an important role in explaining the differences between individual and aggregate wages in Britain.

An additional general problem of aggregation is that it assumes that the relationship between real wages and the business cycle is the same for all individuals or groups of individuals. If wrong, the estimates of real wage cyclicality include a specification bias.

In the last two decades, a number of studies based on micro-panel data for the U.S. found, without exception, robust evidence in favor of a procyclical behavior of real wages.[1] Recently, the focus of micro-data studies on wage cyclicality has been the differences in individual real wage responses to aggregate labor market conditions for different types of workers classified by demographic characteristics, job mobility status, industry, type of payment, etc. Solon *et al.* (1994) and Shin (1994), for example, analyzed the heterogeneity in real wage responses over the cycle between job changers and job stayers. After Bils' (1985) finding that wages are strongly procyclical only among job changers, these authors have reported evidence of considerable real wage procyclicality even for job stayers. Ziliak *et al.* (1999) investigate how real wages respond to local and aggregate unemployment over time for different groups of workers classified according to education, race, industry, occupation and union status. Overall, they found spatial and timing differences in the cyclicality of real wages as well as substantial heterogeneity in wage behavior over the cycle across different groups of workers. Grant (2001) compares the cyclical behavior of union and nonunion wages in the U.S., concluding that substantial reductions in union wage procyclicality since the mid-1980s are associated with reductions in the procyclical exercise of bargaining power. Devereux (2001) provides a detailed analysis of the wage cyclicality of stayers, taking into account the different types of payment. Taken as a whole, his results revealed that incentive pay is much more sensitive to the business cycle than are wage rates or salaries.

The aim of this study is to investigate real wage cyclicality in Portugal for the period 1986-98, addressing the issue of heterogeneity in wage responses to aggregate labor market conditions for workers' hirings and separations. For this purpose a micro-longitudinal panel data obtained from *Quadros de Pessoal (QP)* will be used. *QP* is an annual mandatory employment survey collected by the Portuguese Ministry of Employment. It covers almost all establishments with wage earners. Reported data cover the establishment itself, the firm and each

[1] Meghir and Whitehouse (1996) used data on male employees for the U.K. and also showed that real wages are highly procylical.

of its workers. Since unique identifiers are available for both firms and workers, firms and individuals can be tracked over the sample period. Currently, the data set collects data on about 200,000 firms and 2 million employees.

Our paper is related in at least two distinct ways to the recent microdata studies on wage cyclicality. First, it confirms the empirical evidence of a pro-cyclical behavior of real wages, even for a European country which has a very different labor market institutional framework than the U.S. Second, it addresses the issue of heterogeneity in wage responses to unemployment for workers' accessions and separations, showing that examining wage cyclicality of employed workers without taking into account the flows of workers into and out of the workforce may mask the estimates of real wage cyclicality.

This study will be organized as follows. In the next Section a brief literature review of previous research on real wage behavior over the business cycle is presented. In Section 3 some selected labor market indicators for Portugal are reported. In Section 4 the data set and methodology are described. The empirical framework and main results are discussed in Section 5. Conclusions are outlined in Section 6.

2 Previous Research

In this section a brief description of the empirical literature on the movements of real wages over the cycle will be made. These empirical studies can be classified in two groups, according to whether aggregate or micro-panel data are used. A more exhaustive account of the studies that use micro-panel data will be given. This option is justified by the fact that our own work will use micro-panel data, which will require comparisons with other similar studies. In any case, we do not intend to present an exhaustive and complete survey of the studies published on real wage movements.[2]

2.1 Studies using Aggregate Data

Over the last 60 years economists have been discussing real wage behavior over the cycles. Neoclassical and Keynesian models predicted that real wages behave countercyclically, since an increase in employment can only occur with a decrease in real wages.[3] However, in the presence of imperfect competition, uncertainty, lagged responses, or contracting between firms and workers, there are few reasons to expect this behavior [Brandolini (1995)]. The macroeconomic

[2] For a insightful survey see Brandolini (1995).

[3] Quoting the General Theory [Keynes (1936), p. 17]: " ...with a given organization, equipment and technique, real wages and the volume of output (and hence of employment) are uniquely correlated, so that, in general, an increase in employment can only occur to the accompaniment of a decline in the rate of real wages. Thus I am not disputing this vital fact which the classical economists have (rightly) asserted as indefeasible. ...The real wage earned by a unit of labor has a unique inverse correlation with the volume of employment."

disequilibrium model of Barro and Grossman (1971), for instance, provided a theoretical explanation for the possibility of procyclical real wages. They show that the impact of excess supply of commodities on labor demand removes the one-to-one classical relationship between real wage and employment. In a general disequilibrium situation, unemployment can coexist with "non-excessive" real wages, and a procyclical pattern of real wages is consistent with the theoretical model. On the other hand, implicit contract theory predicts the existence of acyclical real wages [see Rosen (1985) for a recent survey]. According to this theory, contract wages embody implicit payments of insurance premiums by workers in favorable states of nature and receipt indemnities in unfavorable states, meaning that wages are somewhat insulated from current labor market conditions.

The empirical work tried to shed some light on these alternative explanations. In fact, it is worth noting that the establishment of a stylized fact concerning the typical wage pattern over the cycle is purely an empirical matter.

Several studies using aggregate data appeared in the 1970s and 1980s with the aim of analyzing the movements of real wages over the business cycle. The focus of these studies was the raw correlation between real wages and employment or output (or another cyclical indicator). In this sense, real wages are acyclical when they are uncorrelated with the variable chosen as indicator of the business cyce.

Despite the use of appropriate econometric techniques and representative data sets, the conclusions reported in these studies were mixed and inconclusive. Indeed, studies by Chirinko (1980), Canzoneri (1978), Tatom (1980), Neftci (1978), Sargent (1978), and Mehra (1982) provide evidence of countercyclical real wages. However, Bodkin (1969) and Michie (1987) showed that real wages are procyclical. On their side, Geary and Kennan (1982) found evidence for 12 OECD countries in favor of the hypothesis that real wages and employment are statistically independent. Finally, Sumner and Silver (1989) concluded that real wages were either procyclical or countercyclical depending on the sample period chosen.

Table 1.1 displays a selected number of studies on real wage cyclicality based on aggegated data and the main results obtained concerning the correlation between real wages and the cyclical indicator.

Table 1.1: Studies of Real Wage Cyclicality that Use Aggregate Data

Author	Country	Sample	Cyclical Indicator	Real Wage/Cyclical Indicator Correlation	Remarks
Bodkin (1969)	U.S.	Manufacturing data quarterly (annual); 1900-65	Unemployment rate	negative, significant	
	Canada	Manufacturing data quarterly (annual); 1921-65	Unemployment rate	not significant	
Otani (1978)	14 Industrial countries	Manufacturing data annual; 1952-75	Index of industrial production	negative, significant in 6 of the 14 countries	
Chirinko (1980)	U.S.	Manufacturing data annual; 1955-75	Index of industrial production	negative, significant	account for compositional changes in the workforce
Canzoneri (1978)	Canada	Manufacturing data quarterly; 1954-70	Worked hours per capital unit	negative, significant	account for the level of capital stock
Tatom (1980)	U.S.	Private business sector annual; 1948-73	Worked hours per capital unit	negative, significant	account for the cyclical variation of utilization of the capital stock
Neftci (1978)	U.S.	Manufacturing data monthly; 1948-71	Production workers in manufacturing	negative, significant	introduced dynamic analysis
Sargent (1978)	U.S.	Manufacturing data quarterly; 1948-72	Production workers in manufacturing	negative, significant	introduced dynamic analysis

Table 1.1: Continued

Author	Country	Sample	Cyclical Indicator	Real Wage/Cyclical Indicator Correlation	Remarks
Geary and Kennan (1982)	12 OECD countries	Manufacturing data quarterly; 1947-77	Employment in manufacturing	no correlation for the 12 countries[a]	[a]employment and real wages are statistically independent
Sumner and Silver (1989)	U.S.	Manufacturing data annual; 1900-85	Total employment in manufacturing	negative and positive depending on the sample period	
Mehra (1982)	U.S.	Manufacturing data[b] quarterly; 1956-70	Production workers in manufacturing	negative, significant in 9 of the 14 industries	[b] disaggregated by industry at the two digit level
Burda (1985)	U.S.	Manufacturing data[b] annual; 1949-78	Total worked hours	inconclusive	[b] disaggregated by industry at the two digit level
Michie (1987)	U.K.	Manufacturing data[b] annual; 1948-74	Industry output	positive, significant in all industries except one	[b] disaggregated by industry at the two digit level

2.2 Studies using Micro-data

Empirical studies on real wage movements that use micro-panel data began to appear in the mid 1980s. The authors of these studies realized that the changes in the composition of the workforce over the cycle may affect the behavior of aggregate wages, inducing a bias. The usual argument is that less skilled workers are more vulnerable to layoffs in a period of recession. Thus, in periods of business cycle recession real wages are probably averaged over a group of workers with higher wages than those averaged in a period of business cycle expansion, imposing a countercyclical bias. Bils (1985) and Mitchell *et al.* (1985) were the first to explicitly address the effects of using aggregate data on the estimates of real wage cyclicality.

Using two different data sets, the National Longitudinal Survey (NLS) and the Panel Study of Income Dynamics (PSID), these studies point quite decisively toward a procyclical behavior of real wages. Concerning the sign of the composition bias, the results were not so conclusive. Those studies that used the PSID data concluded that aggregation induces a countercyclical bias. The studies using the NLS data showed a relatively unimportant countercyclical composition bias (Bils, 1985) or a procyclical one (Keane *et al.*, 1988). A description of the empirical methodology and main results of these studies will now be presented.

The empirical model usually used in micro-data studies on wage cyclicality is a classic human-capital wage equation with a control for business cycle conditions:

$$W_{it} = \alpha_i + \beta CYC_t + \gamma t + \theta X_{it} + \varepsilon_{it} \tag{1.1}$$

where W_{it} is the natural log of the real wage of individual i in period t, CYC_t is the cyclical indicator, t a time trend and X_{it} a vector of worker characteristics. The parameter β measures the percent wage change in response to a one-point increase in the cyclical indicator. Equation (1.1) may also be specified in growth rate form.

Bils (1985) used the NLS sample of young men and estimated a first differences version of equation (1.1) for the years 1966-80. The NLS is a nationally representative sample of the U.S. male population aged 14 to 24 years drawn in 1966.[4]

The real wage was defined as the average hourly earnings including overtime and deflated by the gross national product (GNP) deflator. The average national unemployment rate was used as the cyclical indicator. The vector of worker characteristics, X, includes experience, education, change in region and in local labor force size, and a dummy for unemployment status during the year. The estimates results reported a β of -1.59 and -1.84 for white men and black men, respectively. He also found that real wages are more procyclical for job changers and for people moving in and out of the workforce, than for job stayers.

In order to have a data set that mimics the data used in aggregate studies, Bils constructed an average aggregate real wage measure by taking the NLS

[4] At that time 5225 individuals were interviewed.

sample and aggregating it by time period. A coefficient estimate for the unemployment rate of -1.66 was obtained. This somewhat surprising result, when compared with past studies based on aggregate data, led Bils to re-estimate the model for the total private workforce and for manufacturing only. A significant and negative relationship between changes in real wages and changes in unemployment was also exhibited by these data. Finally, a regression was estimated for manufacturing workers but excluding overtime pay from the wage measure. In this case the coefficient of changes in the unemployment rate was reduced from -1.08 to -0.65. Thus, Bils concluded that differences from previous studies with aggregate data are partly due to the inclusion, in his work, of overtime pay, and to the difference in sampling period.

Finally, Bils examined the possible biases implicit in aggregating and showed that the average aggregate real wage is countercyclically biased by persons who enter the workforce in times of high employment. In fact, the effect of persons with lower wages joining the employed workforce in good times is to increase the coefficient of the change in unemployment rate by about 0.19 percentage points. This bias, however, is small relative to the strongly procyclical behavior found in the study.

Keane *et al.* (1988) adopted a different methodology. Their major focus is on the sample selection bias that may arise when wages of only employed workers are followed over the cycle. In fact, a bias may arise if individuals who move into and out of the workforce are not randomly chosen from the population with respect to their heterogenous attributes. Using the NLS for the period 1966-81 (excluding again 1972, 1974, 1977 and 1979) they estimated a wage equation like (1.1), where W is an hourly straight-time measure deflated by the consumer price index (CPI), U is the national unemployment rate and X includes education, experience, experience squared and a race dummy. A correction for aggregation bias was made using a self-selection technique. Thus, equation (1.1) was estimated by OLS with no correction for selectivity and by maximum likelihood including a correction for selectivity. In the latter case, an employment equation based on a binary choice model was also estimated by maximum likelihood. The maximum likelihood estimations obtained with no individual effects, with fixed effects and with random-effects models, showed a procyclical real wage behavior. The specification tests, however, favor the latter one and reject the no-effects and fixed effects models. A random effects estimate of -0.66 was obtained for β and a value of -0.96 for the OLS estimate. These results suggest that the selection effect biases the behavior of real wages in a procyclical direction, contrary to Bils' finding of a countercyclical bias. The aggregate results (based on NLS data) also showed that a procyclical bias is present if the selection of workers into and out of the workforce is ignored.

Some robustness tests were made in order to check the previous results. Using the producer price index (PPI) instead of the CPI results in a stronger procyclical estimate of β. The inclusion of overtime pay did not change the unemployment coefficient. The results were also robust to changes in the cyclical indicator (capacity utilization, industrial production and employment indicators).

Using the NLS for young women, Tremblay (1990) investigated the real wage pattern over the cycle for women. The sample includes individuals aged 16 or over and working at least 35 hours a week, corresponding to a total of 4053 observations for white women and 1396 observations for nonwhite women. Equation (1.1) was estimated for the period 1968 to 1978, 1974 and 1976 excluded, using a fixed-effects model. The same model was re-estimated adopting a correction for selectivity.

The cyclical indicator is the local unemployment rate. Education, experience (and its square), seniority (and its square), dummies for residence, union status, marital status, occupational training, industry and occupation are also included as explanatory variables. The dependent variable is the natural logarithm of the hourly wage rate. Parameter estimates on the local unemployment rate are -0.4 and -0.5 for white women and nonwhite women, respectively. The inclusion of a selectivity term did not substantially affect the parameter estimates or t-statistics of the unemployment rate. Estimates of -0.9 for white women and of -0.2 (not statistically significant) for nonwhite women were obtained when the national unemployment rate was used.

The same procedure was applied for NLS data for young men and for the same period (excluding, in this case, 1972, 1974 and 1977). The parameter estimates on the local (national) unemployment rate were considerably higher than those obtained for women, -0.9 (-1.5) for white men and -0.6 (-1.6) for nonwhite men. Some explanations were advanced for the higher procyclicality of men's real wage. The first relies on the fact that male workers are more likely to acquire specific training, which may lead them to opt for wage rather than employment fluctuations over the cycle in order not to lose their previous investments in firm-specific training. Another potential explanation comes from the fact that white men are more heavily concentrated in cyclically sensitive industries such as manufacturing and construction. Finally, the third explanation relies on the differences in labor supply elasticities for men and women.

Shin (1994) extended the NLS-based studies of Bils and Tremblay by including data through 1981 and by examining two types of heterogeneity in real wage cyclicality. He investigated the presence of heterogeneity in real wage cyclicality between job changers and job stayers and between different industries. He also showed that real wages move procyclically, even though wages are undoubtedly more procyclical for changers than for stayers. Regarding heterogeneity by industry, he found that all industries (at the one-digit level) except agriculture exhibit procyclical wage patterns, even though the magnitude of the coefficients varies considerably by industry.

Raisian (1983) investigated the cyclical sensitivity of real weekly wages, weeks and hours worked using the PSID data over the period 1967-79. The sample includes 3828 male heads of household not self-employed and aged 65 or less. Raisian estimated a model like (1.1) using a fixed-effects estimator. The cyclical indicator is the unemployment rate in the industry in which the individual works (or worked). An array of control variables included years in current job (and its square), potential years of work experience (and its square), a dummy for union membership and some dummies for occupational categories.

The results revealed that a 1-percentage point increase in the industry unemployment rate is associated with a 0.65% decline in weekly wages. Estimates for various worker tenure and experience levels showed that wages are more procyclical for more senior employees and more experienced workers. Furthermore, white-collar workers exhibit less procyclical variability in wages than do blue-collar workers. While between union and nonunion white-collar workers there are no significant differences in cyclical wage adjustments, blue-collar nonunion workers exhibit less cyclical wage variability than their union counterparts.

Blank (1990) examined the cyclicality of aggregate real wages using the PSID for the years 1969 to 1982. The sample includes all male household heads aged between 20 and 65 who had reported labor-market earnings. Real wages are defined as average hourly earnings, including overtime and second-job income, deflated by the GNP price deflator. The business cycle was proxied by changes in the gross national product.

The aggregate estimates based on 14 pooled cross-sections and on the assumption of a homogenous labor market exhibited a procyclical behavior of real wages, particularly for white men. This result holds when the sample is restricted to individuals employed for two consecutive years, which seems to indicate that it is not the movement of workers in and out of the workforce that induces aggregate wage procyclicality. This result contradicts the one obtained by Keane *et al.* (1988) of a procyclical bias induced by movements of workers in and out of the workforce. Blank justified this apparent contradiction with differences in the data set and wage measure used. Indeed, the observations included in the NLS data are only on young men, which may be more prone to labor-force participation changes. On the other side, the wage measure in the NLS data is a point-in-time measure rather than an annual one, which may lead to the exclusion of individuals who move in and out of the workforce more often.

The results based on the assumption of a heterogenous labor market showed the existence of significant movements in the workforce over the cycle between the sectors of professional, managerial, and administrative workers, manufacturing workers and nonmanufacturing workers. If wages are correlated with sectoral choice and this is not taken into account in the estimation, a composition bias may also arise. However, the results showed that there is no correlation between those two variables, once a standard set of worker characteristics is controlled for.

Solon *et al.* (1994) showed that the true procyclicality of real wages is obscured in aggregate time series because of a composition bias. Using the earnings data of the Bureau of Labor Statistics (BLS) for the period 1967-87, they obtained an unemployment coefficient estimate of -0.6. In order to show that this modest cyclicality in the aggregate real wages (average hourly earnings including overtime and second jobs, deflated by the GNP deflator) is due to a countercyclical composition bias in the aggregate wage statistics, they made use of longitudinal micro-data from the PSID (1967-87).

The first method used to avoid composition bias was to construct a wage statistic that gives fixed weights to the exact same PSID workers over time.

Doing so, a balanced sample of 355 men aged 25-59 that reported positive labor income and at least 100 hours of work for every year from 1967-87 was obtained. The estimated unemployment coefficient based on this sample reached a value of -1.35 for men (-0.46 for women), thus more than double that obtained from the BLS data. This result remains unaffected when wages are measured by the log of the sample mean instead of the sample mean of the log wage, or deflated by the CPI instead of the GNP deflator.

Although statistically significant, these results may not be representative of the entire workforce. In fact, the small sample size led Solon *et al.* to estimate a model for individual data. Hence, a model like (1.1) with a control for experience was estimated for an unbalanced sample of 64847 observations of 7225 men (50531 observations of 6801 women) aged 16 or over and with no restriction on hours. Even though the increased sample size improves the precision in the estimate of the unemployment coefficient, the respective value changed only slightly to -1.40 for men (-0.53 for women).

In a recent paper, Ziliak *et al.* (1999) investigated how real wages respond to local and aggregate unemployment rates over time. The classic human-capital wage equation (adding a term for the national/local unemployment rate) was estimated by two methods: first differences and fixed effects. The data from the PSID includes 12944 observations of 836 male heads of households (aged at least 20 and less than 45 in 1971) that completed schooling, for the years 1971 through 1990. Annual average hourly earnings deflated by the personal consumption expenditure deflator were used as the wage's measure. Experience was used as a control variable.

Unemployment coefficient estimates of -1.15 from the model in first differences and of -1.31 from the fixed effects model were obtained. The difference between these results and the one of -1.40 reached by Solon *et al.* (1994) may be due to the different sample period used.

When the local unemployment rate is introduced into the model a divergence between the results from the fixed-effects and the first-differences model emerges. This disparity disappears when lagged values of the aggregate and local unemployment rates are included in the model. The results indicate that real wages move procyclically with respect to aggregate and local unemployment, but the response to the former is contemporaneous while to the latter it occurs with a lag. The idea is that the aggregate unemployment rate has a rapid impact on wages because it is common across all local economies, while it takes a long time (because information is more costly to gather) to perceive that a local change has occurred.

Finally, some tests were made to evaluate the heterogeneity in the cyclicality of real wages for various demographic groups. Real wages of union workers are less procyclical to both aggregate and local conditions than those of nonunion workers. Real wages for manufacturing and blue-collar workers are also less procyclical in response to changes in the aggregate unemployment rate. Differences in real wage cyclicality by race were also reported. In particular, for contemporaneous changes in aggregate unemployment the rates of return to education for blacks are countercyclical while for whites are procyclical.

Using data from the PSID, Devereux (2001) investigates the cyclicality of wages within employer-employee matches for the years 1970-91. The sample includes male employees aged between 18 and 64 who worked at least 100 hours in the calendar year. Estimating a first differences version of equation (1.1), a coefficient of -1.16 for the effect of changes in the unemployment rate on changes in average hourly earnings was obtained. The coefficient takes the value of -0.81 for the sample of workers who did not change employers (job stayers). Finally, and because Devereux's main interest is in the cyclicality of wages within employer-employee matches, rather than the cyclicality of wages of workers who do not change employers, he re-estimates the model for the sample of job stayers who do not hold any extra job. With the exclusion of income from extra jobs from average hourly earnings, he gets an unemployment coefficient estimate of -0.54, suggesting that the wage cyclicality of stayers overstates the cyclicality of wages within matches because earnings from extra jobs exert a procyclical bias.

Another important contribution of this study is to examine in detail wage cyclicality for different types of payment. The results indicate that average hourly earnings are much more procyclical for the small group of workers who were paid by piece rate or commissions, than for hourly or salaried workers.

Finally, we shall mention the work of Meghir and Whitehouse (1996), who used data on male employees from the U.K. Family Expenditure Survey, a time series of repeated cross-sections, for the years 1968-86. Using instrumentral variables to control for the possible endogeneity of the unemployment rate, they showed that wages in the U.K. are highly procyclical. A 1 percentage point increase in unemployment leads to a 1.4% fall in hourly real wages. Ignoring the endogeneity of the aggregate unemployment rate yields a smaller coefficient for the unemployment rate (-1.1%).

In Tables 1.2 and 1.3 a synthesis of the methodology and main empirical results of the U.S. micro-data studies is provided. Table 1.2 refers to the studies that used the NLS data and Table 1.3 refers to the studies that used the PSID data.

Table 1.2: Studies of Real Wage Cyclicality that Use Micro-data
National Longitudinal Survey (NLS)

Author	Sample	Model and Estimation Method	Wage Measure	Cyclical Estimates (aggregate unemployment rate)
Bils (1985)	young men	first differences	weekly earnings, including	$\beta = -1.59$ (whites)
	1966-80	(GLS)	overtime, divided by weekly hours	$\beta = -0.64$ (job stayers)
	11 spells		(GNP deflator)	$\beta = -3.69$ (job changers)
Keane *et al.* (1988)	young men	random effects	hourly straight-time earnings	$\beta = -0.96$ (OLS)
	1966-81	(maximum likelihood (ML))	(CPI deflator)	$\beta = -0.66$ (ML)
	12 spells			
Tremblay (1990)	young women	fixed effects	hourly wage	$\beta = -0.90$ (whites)
	1968-78	(OLS)	(CPI deflator)	$\beta = -0.40^*$ (whites)
	9 spells			* local unemployment rate
Shin (1994)	young men	first differences	annual earnings divided by	$\beta = -1.65$ (whites)
	1966-81	(OLS)	annual hours	$\beta = -1.18$ (job stayers)
	12 spells		(GNP deflator)	$\beta = -2.67$ (job changers)

Table 1.3: Studies of Real Wage Cyclicality that Use Micro-data
Panel Study of Income Dynamics (PSID)

Author	Sample	Model and Estimation Method	Wage Measure	Cyclical Estimates (aggregate unemployment rate)
Raisian (1983)	men aged 65 or less 1967-79 13 spells	fixed effects (OLS)	total annual labor income excluding income from extra jobs, divided by weeks worked (CPI deflator)	$\beta = -0.65^*$ * industry unemployment rate
Solon *et al.* (1994)	men/women aged 16 plus 1967-87 21 spells	first differences (OLS)	annual earnings divided by annual hours (GNP deflator)	$\beta = -1.40$ (men) $\beta = -0.53$ (women) $\beta = -1.24$ (men stayers)
Ziliak *et al.* (1999)	men aged 20-65 1971-90 20 spells	fixed effects (FE) first differences (FD) (OLS)	annual earnings divided by annual hours (personal consumption expenditure deflator)	$\beta = -1.31$ (FE) $\beta = -1.15$ (FD)
Devereux (2001)	men aged 18-64 1970-91 22 spells	first differences (OLS)	annual earnings divided by annual hours (personal consumption expenditure deflator)	$\beta = -1.16$ (full sample) $\beta = -0.81$ (job stayers) $\beta = -0.54$ (job stayers with no extra jobs)

3 Some Selected Labor Market Indicators[5]

3.1 Labor Force Participation and Unemployment

Table 1.4 and Figures 1.1 and 1.2 display the evolution of the unemployment and participation rates over the period 1985-98.[6]

The national unemployment rate is obtained from the Labor Force Survey of the *Instituto Nacional de Estatística (INE)*. As can be seen, the period 1985-98 comprises the 1985-92 expansion and the 1993-96 recession. In fact, the unemployment rate reached its peak in 1985 at 8.5%. Since then its evolution has accompanied the economic expansion of the mid-1980s, so that it steadily decreased until 1992. By this time, the Portuguese labor market experienced a situation of near full employment. However, due to the recession initiated in 1993, unemployment rates rose, averaging 7.3% in 1996.

In the mid-1980s and the early 1990s the proportion of female unemployment was much higher than the proportion of male unemployment (see Table 1.4). Since then, male and female unemployment shares have been very similar, albeit slightly higher for women.

Over the 1985-98 period the participation rate seems to exhibit a procyclical behavior. The evolution of male activity rates over this period is mainly the result of a declining participation of younger (less than 24 years) and older (50-64 years) men, and an almost unchanged participation behavior for men in the 24-50 years group. In its turn, the rise in female participation is the result of a long-term trend mainly explained by changing life styles and family strategies concerning participation in the labor market. Portugal exhibits one of the highest rate of female participation in the labor force in the European Union.

[5] In the Annex a brief description of the Portuguese labor market is presented.

[6] It should be noted that in 1992 and 1998 there is a series break, implying that the results are not directly comparable with the previous year.

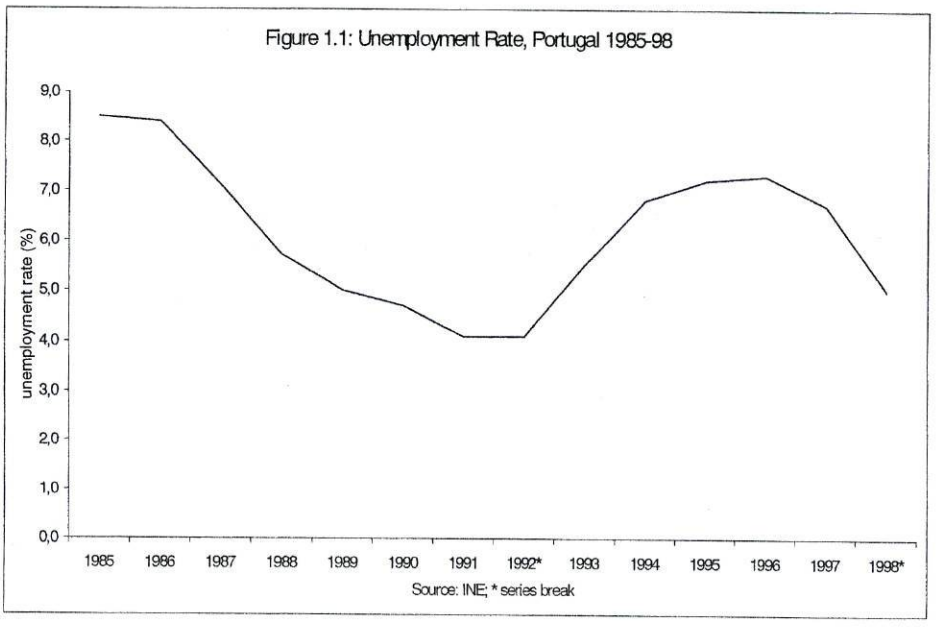
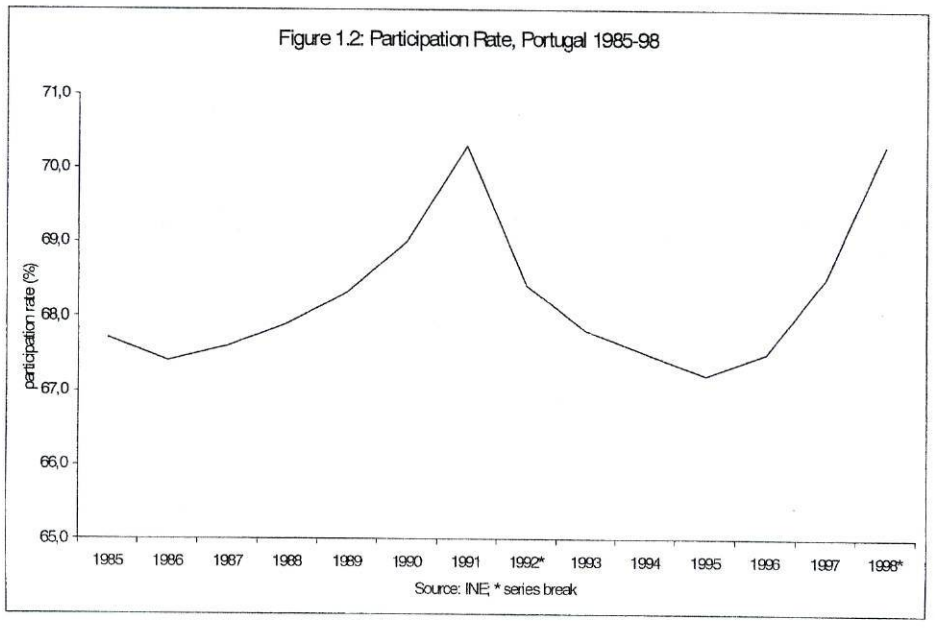


Table 1.4: Unemployment and Participation Rates
Portugal, 1985-98

	Unemployment Rate (%)			Participation Rate (%)[a]		
Year	All	Male	Female	All	Male	Female
1985	8.5	6.3	11.4	67.7		
1986	8.4	6.5	11.1	67.4		
1987	7.1	5.4	9.5	67.6	80.0	55.6
1988	5.7	4.1	8.0	67.9	79.9	56.7
1989	5.0	3.4	7.1	68.3	80.4	57.2
1990	4.7	3.2	6.6	69.0	80.5	58.2
1991	4.1	2.8	5.8	70.3	80.7	60.6
1992*	4.1	3.5	4.9	68.4	78.7	58.9
1993	5.5	4.7	6.5	67.8	77.2	59.0
1994	6.8	6.0	7.8	67.5	76.4	59.3
1995	7.2	6.4	8.0	67.2	75.4	59.4
1996	7.3	6.5	8.2	67.5	75.5	59.9
1997	6.7	6.1	7.6	68.5	76.4	61.1
1998*	5.0	4.0	6.2	70.3	78.6	62.4

Source: INE; * series break.

[a] Men and women aged 15-64 years.



3.2 Real and Nominal Wages

In Tables 1.5 and 1.6 the evolution of nominal and real wages over the 1985-98 period is presented, using data from *Quadros de Pessoal* (*QP*). The *QP* data set includes all dependent workers, excluding public administration and non-market services.[7] The wages were deflated by the CPI (base=1985).

Table 1.5 provides information on annual rates of growth of average monthly base wages and average monthly earnings.[8] Portuguese nominal wages and earnings show an upward trend over the entire period 1985-98. The rates of growth, however, have decreased in recent years, keeping pace with inflation. In fact, in the 1990s a great effort was made by the government and social partners in order to reduce inflation and obey the convergence criteria of the Euro area (the figures for the inflation rate are shown in column 1 of Table 1.7).

According to Table 1.6, changes in real wages were also positive, except for the year of 1995. Whereas the 1989 fall in the rate of growth of real earnings was mainly attributable to a transitory failure in expectations following the setback in disinflationary policies, the 1994/95 decline stems from the recession that took place in 1993.

Table 1.7 provides the rates of change of nominal and real earnings for the whole economy. The inflation rate as measured by the CPI (excluding rent) for the mainland is also reported. The evidence for the entire economy displays essentially the same pattern.

[7] A detailed description of this data set will be presented in the next section.

[8] The average monthly earnings include base wages, seniority payments and regular benefits.

Table 1.5: Nominal Monthly Base Wages and Earnings
Portugal, 1985-98

Year	Average Monthly Base Wages*	Percentage Change	Average Monthly Earnings*	Percentage Change
1985	24719.4		27526.2	
1986	29811.7	20.6	33239.9	20.8
1987	34242.3	14.9	38332.2	15.3
1988	37550.3	9.7	42341.8	10.5
1989	42582.4	13.4	47955.3	13.3
1990	n. a.	n. a.	n. a.	n. a.
1991	58179.9	n. a.	66319.4	n. a.
1992	66439.7	14.2	76985.1	16.1
1993	73797.4	11.1	84944.2	10.3
1994	80192.8	8.7	92943.5	9.4
1995	82518.4	2.9	95890.5	3.2
1996	87169.6	5.6	101744.6	6.1
1997	90557.7	3.9	106014.9	4.2
1998	94470.0	4.3	110825.1	4.5

Source: *QP*; 1986-93 March data; 1994-98 October data; n. a.: not available.
* in PTE (escudo): 1 EUR≡200.482 PTE.

Table 1.6: Real Monthly Base Wages and Earnings
Portugal, 1985-98

Year	Average Monthly Base Wages*	Percentage Change	Average Monthly Earnings*	Percentage Change
1985	24719.4		27526.2	
1986	26563.9	7.5	29631.2	7.6
1987	27786.8	4.6	31114.8	5.0
1988	28167.7	1.4	31756.3	2.1
1989	28428.3	0.9	32017.0	0.8
1990	n. a.	n. a.	n. a.	n. a.
1991	30693.8	n. a.	34984.1	n. a.
1992	32297.7	5.2	37409.9	7.0
1993	33420.3	3.5	38472.5	2.8
1994	33761.2	1.0	39134.1	1.7
1995	33360.2	-1.2	38773.2	-0.9
1996	34222.3	2.6	39956.1	3.0
1997	34924.0	2.0	40878.3	2.4
1998	35305.3	1.1	41417.6	1.3

Source: *QP*; 1986-93 March data; 1994-98 October data; n. a.: not available.
* in PTE (escudo): 1 EUR≡200.482 PTE.

Table 1.7: Percentage Changes of Prices and Earnings
Portugal, 1985-98

Year	Consumer Price Index[1] Annual Percentage Change	Compensation per employee*[2]	
		Nominal Percentage Change	Real Percentage Change
1985	19.3		
1986	11.7		
1987	9.4	13.8	3.8
1988	9.6	11.6	0.7
1989	12.6	14.1	1.2
1990	13.4	17.6	4.9
1991	11.4	19.1	6.6
1992	8.9	13.8	3.9
1993	6.5	8.0	0.8
1994	5.2	5.5	-0.1
1995	4.1	7.1	2.9
1996	3.1	6.5	3.3
1997	2.2	5.8	3.5
1998	2.7	5.8	2.8

Sources: [1] INE; [2] Banco de Portugal (1999), *Annual Report 1998.*

* Estimates for the entire economy, annual basis. Employers' contributions to Social Security included. Private consumption deflator.

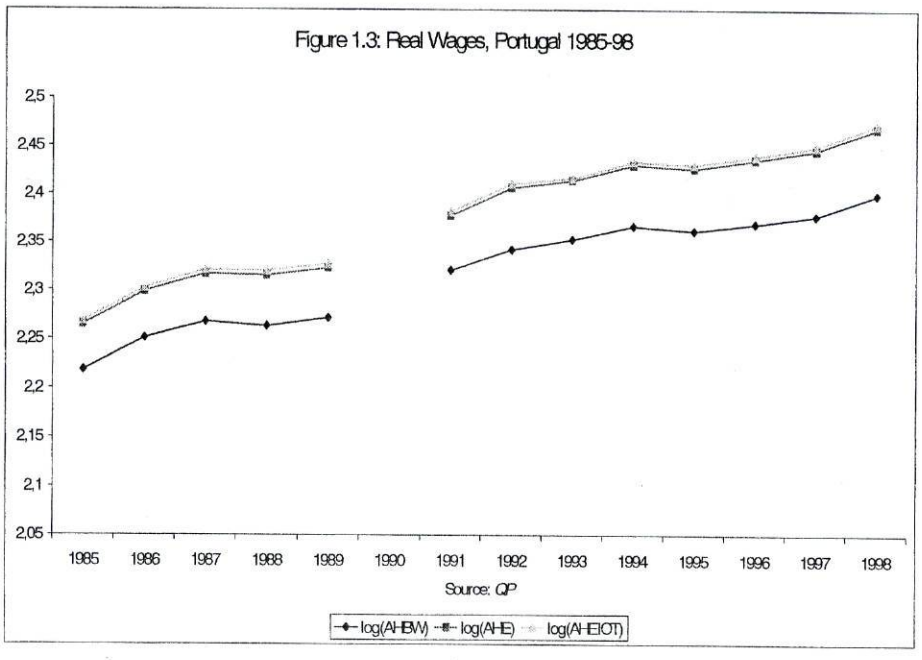
The standard measure of wages used in studies of the cyclical behavior of real wages is an hourly measure. Figure 1.3 displays the evolution of real hourly wages and earnings over the period 1985-98 using the *QP* data. Three measures of hourly wages were defined: average hourly base wages (AHBW), average hourly earnings (AHE) and average hourly earnings including overtime pay (AHEIOT).[9]

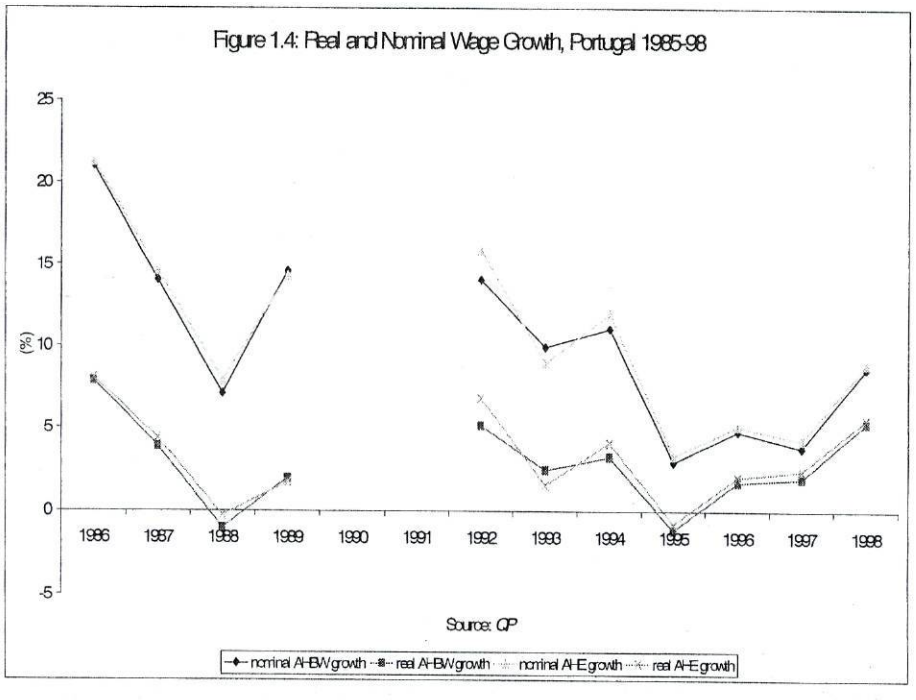
In real terms, average hourly base wages and average hourly earnings (with or without overtime pay included) exhibit a reasonably steady increase in the 1985-98 period. Between 1985 and 1998, AHE grew around 60% in real terms, which corresponds to an average annual growth rate of 4% (3.5% for hourly AHBW). The same is true when overtime pay is included. In fact, average hourly earnings including overtime pay (AHEIOT) present an evolution pattern quite similar to AHE. This is mainly due to the fact that in this period overtime hours represent, on average, only 0.9% of total hours worked, giving rise to an almost coincidence of values between AHEIOT and AHE.

Figure 1.4 displays the rates of growth of real and nominal hourly base wages and earnings. Independently of the wage measure used (base wages or earnings), real and nominal wage growth rates present a similar pattern over the period 1985-98.

However, when these results are compared to those obtained for monthly base wages and earnings (see Table 1.6), some differences emerge in the evolution of the rates of growth of real wages and earnings regardless of whether an hourly or monthly measure is used. In 1994, for example, even though the rate of growth of real monthly wages and earnings decreased, a reduction in the total number of usual hours worked contributed to a higher growth in real hourly wages/earnings.

[9] AHE equals the ratio of total regular payroll to total normal hours and AHEIOT is defined as the ratio between total regular payroll including overtime pay and the sum of normal and extra hours of work. Total regular payroll includes base wages, seniority payments and regular benefits.


Figure 1.3: Real Wages, Portugal 1985-98
2,5
2,45
2,4
2,35
2,3
2,25
2,2
2,15
2,1
2,05
1985 1986 1987 1988 1989 1990 1991 1992 1993 1994 1995 1996 1997 1998
Source: QP
log(AHEW) log(AHE) log(AHEOT)


Figure 1.4: Real and Nominal Wage Growth, Portugal 1985-98
25
20
15
10
5
0
-5
(%)
1986 1987 1988 1989 1990 1991 1992 1993 1994 1995 1996 1997 1998
Source: QP
nominal AHEW growth real AHEW growth nominal AHE growth real AHE growth

4 Data and Methodology

4.1 Data Description

The data sets of this study were constructed using the data from *Quadros de Pessoal* (*QP*). *QP* is an annual mandatory employment survey collected by the Portuguese Ministry of Employment, that covers virtually all establishments with wage earners.[10] Indeed, each year every establishment with wage earners is legally obliged to fill in a standardized questionnaire. Reported data cover the establishment itself (location, economic activity and employment), the firm (location, economic activity, employment, sales and legal framework) and each of its workers (gender, age, education, skill, occupation, tenure, earnings and duration of work). The information on earnings is very complete. It includes the base wage (gross pay for normal hours of work), seniority payments, regular benefits, irregular benefits and overtime pay, as well as the mechanism of wage bargaining. Information on normal and overtime hours of work is also available.

Twelve spells of *QP*, from 1986 to 1998, were available for this study.[11] From 1986 to 1993 the information was collected in March of each year, and since 1994, in October.

There are three main reasons that make this survey a good source for the study of wage cyclicality. The first is its coverage. By law, the questionnaire is made available to every worker in a public space of the establishment. This requirement facilitates the work of the services of the Ministry of Employment that monitor compliance of firms with the law (e. g., illegal work). Indeed, the administrative nature of the data and its public availability implies a high degree of coverage and reliability. Second, this survey is conducted on a yearly basis, and its identifying scheme allows accurate identification of workers making it possible to track them over the years. The workers' identification number is based on a transformation of his/her social security number. Finally, this source enables the matching of firms and its workers, which allows us to classify the situation of the worker on the job (stayer/changer, accession/separation). Each firm entering the database is assigned a unique identifying number and the Ministry implements several checks to ensure that a firm that has already reported to the database is not assigned a different identification number.

Naturally, this data source also has its own limitations. The most important one for the purpose of analyzing wage cyclicality is the short time period covered (1986-98, excluding 1990).

Table 1.8 reports information on the number of firms and workers covered by *QP* in each year of the 1986-98 period.

[10] Public administration and non-market services are excluded.

[11] No computer files are available for the year 1990.

Table 1.8: Data set Coverage
Quadros de Pessoal, 1986-98

Year	Number of Firms	Number of Workers
1986	106770	1897785
1987	110865	1936801
1988	122775	1996802
1989	137155	2169835
1990	n.a.	n.a.
1991	148602	2233237
1992	159192	2268151
1993	165876	2215481
1994	184306	2202609
1995	192270	2232548
1996	197558	2233721
1997	213589	2350782
1998	224456	2430691

Note: n. a. denotes not available.

4.2 Methodology

As explained above, micro-data allow us to avoid the consequences of composition bias that might contaminate the results when aggregate data are used. Thus, in order to decompose the effects of changes in the composition of the workforce over the cycle and its effects on real wage cyclicality, four different samples of individuals will be analyzed in this study.

The first includes all the individuals that are present in every year of the twelve years surveyed. Since unique identifiers are available for each worker, individuals can be followed over the entire sample period. A separate analysis will also be made for those workers that never changed employer over all that period (job stayers). Even though this method of analyzing the exact same workers from the *QP* files over the time avoids the problem of composition bias, it does not avoid the problem of selectivity bias that may arise when wages of ever employed workers are followed over the cycle. In fact, the population represented by this sample may experience different wage cyclicality than is experienced by other groups in the labor force, namely those with a lower attachment to employment/workforce. In order to surpass these selectivity problems, a second data set that contains the individuals employed for two consecutive years was also constructed.

Finally, and in order to evaluate heterogeneity in wage responses to aggregate unemployment between continuously employed workers and hirings and separations, two additional data sets were constructed. One that includes newly hired workers (accessions) and the other that contains recently separated workers (separations). Next we will describe more accurately each of these four samples.

As mentioned above, the first data set contains all the workers that are present in each and every one of the twelve years surveyed.[12] This sample was restricted to individuals aged between 18 and 52 years old in 1986, excluding agricultural workers.[13] We have also excluded those individuals for whom some explanatory variable is not available for a particular year, namely those with no information on wages. In order to minimize the effects of outliers in wages, for each wage measure used we dropped 0.5% of the observations corresponding to the top and bottom tails of the wage distribution. After these exclusions, a balanced panel of 39,284 men and a balanced panel of 12,926 women was assembled. Hereinafter, we will call this sample 'long-term employee'. The information on the identification number of the firm allows us to obtain from this data set a subsample of job stayers, i. e, workers that stayed with the same employer over the entire period 1986-98 (long-term employee/job stayers). This subsample is comprised of a balanced panel of 23,809 men and 8,623 women.

The second data set contains a sample of individuals that are present in two consecutive years.[14] This sample was also restricted to individuals aged between 18 and 64 years old, excluding agricultural workers. Those individuals for whom some explanatory variable is not available for a particular year, as well as the wages' outliers, were both excluded from the sample. An unbalanced panel of 170,414 observations on men and 101,750 on women was obtained. We will call this data set 'two-years employee'.

The third data set (hereinafter 'accessions') includes a random sample of individuals that in each year are classified in the *QP* data files as a newly hired worker. A worker is classified in each year as newly hired if his tenure in that year is less or equal to one year. We obtained for the 12 pooled cross-sections, 115,009 observations on men and 72,662 on women.[15]

Finally, the fourth data set (hereinafter 'separations') includes a random sample of individuals that in each year are classified as a recently separated worker. A worker is classified in each year as recently separated if he (she) is present in the *QP* registers in year t but is not present in year $t+1$. In order to control for separations we need to exclude the years of 1989 and 1998. Thus, this sample is comprised of 10 pooled cross-sections corresponding to 88,994 observations on men and 56,979 on women.[16]

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[12] When a worker is present in the *QP* files for more than one time in a given year, the register in the firm in which he had worked a higher number of hours was selected.

[13] In agriculture a considerable amount of payments are non-pecuniary. We thought it better to exclude these workers from the analysis. In any case, the number of these workers is almost negligible.

[14] The sample was drawn according to a normal random number generator.

[15] The same restrictions applied to the sample of two-year employee hold for this sample.

[16] The same restrictions applied to the sample of long-term employee hold for this sample.

5 Empirical Results

5.1 Empirical Model

The empirical model that will be used to test for real wage cyclicality is a classic human-capital wage equation with a control for business cycle conditions. The static form of the model is:

$$W_{it} = \alpha_i + \gamma_1 t + \gamma_2 t^2 + \gamma_3 U_{t-1} + \gamma_4 Z_i + \gamma_5 X_{it} + \gamma_6 X_{it}^2 + \varepsilon_{it},$$

$$i = 1, ..., N; \qquad t = 1, ..., T \tag{1.2}$$

where W_{it} is the natural log of the real wage of individual i in time t, U_{t-1} is the aggregate unemployment rate in period $t-1$, t and t^2 are a time trend and its square, X_{it} is a vector of time-varying worker characteristics such as experience and its square (X_{it}^2), Z_i is a vector of time-invariant worker characteristics such as education, α_i is a vector of unobserved individual-specific characteristics that are fixed over time and ε_{it} is a zero-mean random term with constant variance.

If α_i is correlated with the regressors, estimating (1.2) by ordinary least squares (OLS) would yield biased estimates of the γ's parameters. Whenever possible, a fixed effects estimator (within-group estimator) will be used in order to deal with this potential problem (and, thus, estimate consistently the model parameters). In this case, OLS is applied to the following transformed model of equation (1.2):

$$\widetilde{w}_{it} = \delta_1 \widetilde{t} + \delta_2 \widetilde{t^2} + \delta_3 \widetilde{U}_{t-1} + \delta_4 \widetilde{X}_{it} + \delta_5 \widetilde{X}_{it}^2 + \widetilde{\varepsilon}_{it} \tag{1.3}$$

where $\widetilde{w}_{it}, \widetilde{t}, \widetilde{t^2}, \widetilde{U}_{t-1}, \widetilde{X}_{it}, \widetilde{X}_{it}^2$ and $\widetilde{\varepsilon}_{it}$ are the variables listed in their deviations from individual time-series means. The δ coefficients are defined as: $\delta_1 = \gamma_1 + \gamma_{51}$, $\delta_2 = \gamma_2$, $\delta_3 = \gamma_3, \delta_4 = \gamma_{5j}$ $(j = 2, ..., k)$ and $\delta_5 = \gamma_6$. The parameter γ_{51} refers to the coefficient on experience and the parameters γ_{5j} refer to the coefficients on other time-varying worker characteristics. With this transformation the disturbance term, $\widetilde{\varepsilon}_{it}$, is uncorrelated with the regressors. In fact, in the fixed effects model, unlike the random effects model, consistency of the coefficient estimates may be retained even when the individual-specific terms are correlated with the regressors.

The parameter δ_3 (γ_3) measures the percent wage change in response to a one-point increase in the unemployment rate. A negative value of δ_3 (γ_3) implies that wages rise when unemployment diminishes, so that wages are procyclical. If, on the contrary, δ_3 (γ_3) is positive, wages are countercyclical.

Model (1.2) cannot be applied to the samples of hirings and separations since, by construction, panel data are not available in these two cases. To deal with this problem, equation (1.2) will be estimated by OLS with no control for individual unobserved heterogeneity. In order to enable the comparison

of the results obtained for continuously employed workers and workers' hirings and separations, the OLS estimates of equation (1.2) without individual-specific effects will also be reported for the former.

5.2 Results

In this section the main empirical results obtained from the different regression models are presented. In Section 5.2.1 the OLS and fixed-effects regression results are reported for continuously employed workers. In Section 5.2.2 the OLS regression estimates are shown for workers' hirings and separations. All results are presented separately for men and women.

5.2.1 Continuously Employed Workers

The empirical analysis starts with the study of real wage cyclicality of those individuals that remained employed over the entire period of 1986-98 (long-term employee).

In Table 1.9 the OLS estimates of equation (1.2) with no individual specific effects are presented for two different specifications (men and women separately). In both cases the dependent variable is the natural logarithm of the average hourly earnings (AHE) deflated by the CPI (1985 PTE).[17] The AHE is obtained by dividing the total regular payroll in the month by the total number of normal hours worked.[18] Specification (1) includes a constant (CONST), the unemployment rate (U), a trend (T) and its square (TSQ). Since wages are set at least six months to one year in advance, there is a delayed relationship between wages and economic growth. To capture this lagged effect we use the unemployment rate of the previous year. specification 2 also includes some controls for workers' observed heterogeneity. Education (EDUC) refers to the total number of years of schooling completed.[19] Worker's age is used as a proxy for experience (AGE). Eight dummies that characterize worker's skill are also added.[20] Descriptive statistics (mean, standard deviation, minimum and maximum) of all these variables are provided in Tables A.1.1 (for men) and A.1.2 (for women) of Appendix A.

[17] Between 1986-93 the price index is calculated from March of year $t-1$ to March of year t, whereas in the 1994-98 period the price index is calculated from October of year $t-1$ to October of year t. Between 1993 and 1994 there is an adjustment of a year and a half because in 1993 wages refer to March and in 1994 wages refer to October.

[18] Total regular payroll includes base wages, seniority payments and regular benefits.

[19] It should be noted that some workers obtained a higher educational level while employed.

[20] See Table A.1.1 in Appendix A for a definition of the qualification levels.

Table 1.9: OLS Results
Long-term Employee
(Dependent variable: log Average Hourly Earnings)

	Men		Women	
Independent Variables	(1)	(2)	(1)	(2)
U	-0.98	-1.00	-0.69	-0.75
	(-12.6)	(-15.9)	(-4.3)	(-7.1)
T	0.0328	0.0077	0.0309	0.0006*
	(31.2)	(9.0)	(14.1)	(0.4)
TSQ	0.0005	0.0012	0.0001*	0.0008
	(5.5)	(17.2)	(0.7)	(6.5)
AGE		0.0613		0.0519
		(121.7)		(63.9)
AGESQ		-0.0006		-0.0004
		(-102.2)		(-45.0)
EDUC		0.0594		0.0921
		(288.8)		(283.2)
Qualification Level				
Q1		0.4788		0.3701
		(69.3)		(32.4)
Q2		0.4922		0.4834
		(78.9)		(50.6)
Q3		0.2483		0.2454
		(43.6)		(28.4)
Q4		0.2736		0.2928
		(47.1)		(34.8)
Q5		0.1071		0.0763
		(19.5)		(9.7)
Q6		0.0705		-0.0418
		(12.7)		(-5.3)
Q7		-0.0523		0.0453
		(-8.9)		(5.5)
CONST	5.501	3.685	5.275	3.473
	(847.6)	(299.1)	(389.6)	(183.6)
$\overline{R}^2$	0.12	0.42	0.06	0.61
Sum of squared residuals	77782.7	51144.0	36647.7	15338.1
Number of Observations	471408		155112	

Notes: (i) t-statistics are in parentheses;
(ii) all estimates are significant at 1%, except those with an *.

The OLS estimates show evidence of a moderate procyclical behavior of real wages for both men and women. According to specification 2 of Table 1.9, a 1-percentage point decrease in the national unemployment rate raises average hourly earnings by 1.0 percent for men and by 0.75 percent for women. Controlling for worker observed heterogeneity gives more precise estimates of the unemployment rate coefficient. The less procyclical behavior of women's real wages compared to men's real wages is in accordance with previous findings shown by Tremblay (1990) and Solon *et al.* (1994).

Table 1.10 shows the estimates of the unemployment rate coefficient for specification 2 using three other wage measures. In column 2 the hourly earnings including overtime pay (AHEIOT)[21] measure is used and in columns 3 and 4 the hourly base wage (AHBW)[22] and the monthly base wage (AMBW) are adopted, respectively. For comparison reasons the unemployment coefficient estimates for AHE are reported in column 1.

The inclusion of overtime pay does not change the hourly wage cyclicality of either men or women. Hourly earnings (AHEIOT) decrease by 1.04 percent and 0.70 percent in response to a one percentage point increase in the unemployment rate, for men and women, respectively. This result is not surprising since in the Portuguese labor market overtime hours represent a non-significant percentage of total hours worked. In fact, for the period 1986-98, overtime work for firms employing paid labor corresponds, on average, to 0.9% of the total number of hours worked.

Comparing the figures obtained for the estimates of the unemployment coefficient using an hourly measure (AHBW) or instead, a monthly measure (AMBW), leads us to conclude that hourly wages are more procyclical than monthly wages. According to Table 1.10, a one percentage point decrease in the unemployment rate raises average hourly base wages of male (female) workers by 1.10% (1.08%) and average monthly base wages by only 0.74% (0.28%). This result appears to suggest that, if anything, normal hours worked tend to behave in a slightly countercyclical fashion.[23]

[21] The AHEIOT is defined as the ratio between total regular payroll including overtime pay and the sum of normal and extra hours of work.

[22] The AHBW is defined as the ratio between the monthly base wage and the total number of normal hours worked in the month.

[23] This somewhat unexpected behavior in usual hours of work may be justified by the fact that the period under analysis (mainly an expansionary period) is characterized by a gradual reduction in the average duration of work that reached its peaks in 1991 and 1997 (see *INE - Employment Survey*). In fact, in 1991 a reduction in the normal period of work from 48 to 44 hours per week was legally imposed and in 1996 a further reduction from 44 to 40 hours per week was implemented. This may justify why hourly wages over that period display a more procyclical behavior than wages per employee.

Table 1.10: OLS Results (specification 2)
Long-term Employee
Alternative Wage Measures

	Wage Measure			
	AHE	AHEIOT	AHBW	AMBW
Men				
Cycle Regressor (U)	-1.00	-1.04	-1.10	-0.74
	(-15.9)	(-16.3)	(-21.0)	(-15.2)
Women				
Cycle Regressor (U)	-0.75	-0.70	-1.08	-0.28
	(-7.1)	(-6.6)	(-11.9)	(-3.4)

Notes: (i) t-statistics are in parentheses;
(ii) all estimates are significant at 1%.

In tables 1.11 and 1.12 the same regression results are presented for the subsample of job stayers.[24] The results reported for male and female job stayers exhibit a slightly more procyclical behavior of wages in contrast with previous empirical findings of a higher procyclicality of real wages for job changers [see, for example, Bils (1985) and Shin (1994)]. Once again, the results reveal that monthly wages are less procyclical than hourly wages (see Table 1.12).

[24]Tables A.1.3 (for men) and A.1.4 (for women) in Appendix A contain the descriptive statistics for these samples of workers.

Table 1.11: OLS Results
Long-term Employee/Job Stayers
(Dependent variable: log Average Hourly Earnings)

	Men		Women	
Independent Variables	(1)	(2)	(1)	(2)
U	-1.11	-1.20	-0.79	-0.85
	(-11.5)	(-15.8)	(-4.0)	(-6.7)
T	0.0272	0.0033	0.0280	-0.00008*
	(20.8)	(3.2)	(10.4)	(-0.0)
TSQ	0.0009	0.0016	0.0003*	0.0009
	(8.5)	(18.5)	(1.5)	(6.5)
AGE		0.0549		0.0485
		(91.0)		(50.1)
AGESQ		-0.0005		-0.0004
		(-74.4)		(-35.4)
EDUC		0.0634		0.0933
		(246.9)		(231.4)
Qualification Level				
Q1		0.4859		0.3413
		(53.3)		(23.5)
Q2		0.5253		0.4639
		(66.6)		(39.6)
Q3		0.2643		0.2153
		(36.2)		(19.8)
Q4		0.2782		0.2723
		(37.3)		(25.8)
Q5		0.1125		0.0458
		(15.9)		(4.7)
Q6		0.0117*		-0.0860
		(1.6)		(-8.7)
Q7		-0.0325		0.0228
		(-4.3)		(2.2)
CONST	5.482	3.771	5.280	3.583
	(678.8)	(250.2)	(318.4)	(155.2)
$\overline{R}^2$	0.12	0.46	0.06	0.62
Sum of squared residuals	44246.4	27260.9	24459.4	9991.6
Number of Observations	285708		103476	

Notes: (i) t-statistics are in parentheses;
(ii) all estimates are significant at 1%, except those with an *.

Table 1.12: OLS Results (specification 2)
Long-term Employee/Job Stayers
Alternative Wage Measures

	Wage Measure			
	AHE	AHEIOT	AHBW	AMBW
Men				
Cycle Regressor (U)	-1.20	-1.23	-1.23	-0.83
	(-15.8)	(-16.0)	(-19.2)	(-13.5)
Women				
Cycle Regressor (U)	-0.85	-0.82	-1.16	-0.39
	(-6.7)	(-6.4)	(-10.4)	(-3.8)

Notes: (i) t-statistics are in parentheses;
(ii) all estimates are significant at 1%.

The results obtained from the sample of long-term employee should be interpreted with caution, due to the potential problem of selection bias. In fact, a sample selection bias may arise when only employed workers are followed over the cycle. The bias may arise if those workers who move in and out of the workforce over the cycle have (un)observed characteristics systematically different from those who stay in. Beyond that, the population represented by this sample may experience different wage cyclicality than what is experienced by other groups in the labor force, namely those with more unstable labor market attachments.

In order to avoid these problems, an identical empirical analysis will be made for the sample of workers employed for two consecutive years (two-year employee).[25] Tables 1.13 and 1.14 present the OLS results for both men and women. For male workers, an unemployment coefficient estimate of -1.16 was obtained for specification 2 (see column 2 of Table 1.13), indicating a more procyclical behavior of men's real wages when compared to the unemployment rate estimate obtained for the sample of long-term employee (-1.00). For women, the results for the sample of two-year employee reveal a less procyclical behavior of real wages. An unemployment coefficient estimate of -0.45 (see column 4 of Table 1.13) was obtained against a value of -0.75 for long-term employees. When alternative wage measures are used this same pattern of results holds for both men and women (see Table 1.14).

Concerning the estimates obtained for the regression coefficients of the other explanatory variables, the estimation results reveal that, in general, qualification levels affect wages more significantly for the sample of two-year employee than for the sample of long-term employee. For men, the effect of age is higher for the sample of long-term employee whereas for education the effect is more pronounced for the sample of two-year employee. For women the effect of age and education on wage determination is more pronounced for the sample of long-term employee than for the sample of two-year employee.

[25] Tables A.1.5 (for men) and A.1.6 (for women) in Appendix A provide the descriptive statistics for these samples of workers.

Table 1.13: OLS Results
Two-year Employee
(Dependent variable: log Average Hourly Earnings)

	Men		Women	
Independent Variables	(1)	(2)	(1)	(2)
U	-1.38	-1.16	-0.13*	-0.45
	(-8.3)	(-9.7)	(-0.7)	(-3.3)
T	0.0322	0.0234	0.0294	0.0152
	(12.7)	(12.7)	(10.2)	(7.4)
TSQ	-0.0002*	-0.0004	0.00005*	-0.0001*
	(-0.8)	(-2.9)	(0.2)	(-0.7)
AGE		0.0539		0.0345
		(90.8)		(49.3)
AGESQ		-0.0005		-0.0003
		(-70.8)		(-31.5)
EDUC		0.0639		0.0683
		(177.7)		(163.1)
Qualification Level				
Q1		0.6925		0.7333
		(92.0)		(71.4)
Q2		0.6903		0.6698
		(94.9)		(70.9)
Q3		0.4423		0.4576
		(74.4)		(56.3)
Q4		0.4518		0.5054
		(74.8)		(76.9)
Q5		0.2043		0.1854
		(42.8)		(41.3)
Q6		0.1513		0.1052
		(29.7)		(22.7)
Q7		0.0308		0.0927
		(5.7)		(17.6)
CONST	5.374	3.581	5.024	3.728
	(349.9)	(229.9)	(285.6)	(212.7)
$\overline{R}^2$	0.05	0.50	0.06	0.51
Sum of squared residuals	43732.6	23026.2	21699.1	11082.9
Number of Observations	170414		101750	

Notes: (i) t-statistics are in parentheses;
(ii) all estimates are significant at 1%, except those with an *.

Table 1.14: OLS Results (specification 2)
Two-year Employee
Alternative Wage Measures

	Wage Measure			
	AHE	AHEIOT	AHBW	AMBW
Men				
Cycle Regressor (U)	-1.16	-1.18	-1.03	-0.58
	(-9.7)	(-9.7)	(-10.0)	(-5.2)
Women				
Cycle Regressor (U)	-0.45	-0.41	-0.67	-0.32*
	(-3.3)	(-3.0)	(-5.5)	(-2.2)

Notes: (i) t-statistics are in parentheses;
(ii) all estimates are significant at 1%, except those with an *.

Finally, in order to account for unobserved individual heterogeneity, equation (1.2) was re-estimated using a fixed effects regression model. The results for both data sets and for specification 2 are reported in Table 1.15. The fixed effects estimates of the unemployment rate coefficient obtained for the sample of long-term employee are identical to those obtained for this sample of workers using OLS. The main difference is the precision of the regression coefficients as indicated by the t-statistics, which is visibly higher for the estimation with individual fixed effects.[26] As would be expected, the effects on wages of those variables that account for observed worker heterogeneity are estimated with some trepidation when the fixed effects estimator is used. In fact, in the presence of a fixed effects model, those variables that control for worker heterogeneity are only taken into account whenever a within-individual change occurs, which may magnify the bias induced by measurement error in such variables.

Regarding the sample of two-year employee the OLS and fixed effects estimates diverge considerably in magnitude and precision. For men, the unemployment rate estimate is -0.0045, whereas for women this same coefficient is not statistically significantly different from zero (see columns 2 and 4 of Table 1.15).

In a nutshell, the fixed effects results showed a procyclical behavior of real hourly wages for men, whereas for women the results are not conclusive. The OLS results exhibited a strongly procyclical behavior of men's real hourly wages (including job stayers) and a moderate procyclical behavior of women's real hourly wages.

The comparison of the OLS with the fixed effects regression results for these two samples seems to suggest that unobserved heterogeneity plays a non-trivial role in the cyclicality of wages. Leaving aside cross-sectional heterogeneity is relevant for the sample of two-year employee, but not in the case of the long-term employee. We tentatively speculate that unobserved heterogeneity tends to affect the composition of the workforce in a way that magnifies the cyclicality of real wages.

[26] The same is true for the sub-sample of job stayers. Results available upon request.

Table 1.15: Fixed Effects Results
(Dependent variable: log Average Hourly Earnings)

	Men		Women	
Independent Variables	Long-term Employee	Two-year Employee	Long-term Employee	Two-year Employee
U	-0.98	-0.45	-0.71	-0.04*
	(-34.5)	(-4.6)	(-15.5)	(-0.3)
T	0.0510	0.0340	0.0394	0.0210
	(105.1)	(21.8)	(51.5)	(10.8)
TSQ	0.0008	-0.0001*	0.0004	0.0004
	(24.6)	(-0.5)	(7.4)	(2.6)
AGESQ	-0.0003	0.00005	-0.0002	0.00007
	(-69.8)	(11.2)	(-25.5)	(11.3)
EDUC	0.0076	0.0462	0.0130	0.0370
	(18.7)	(67.0)	(19.8)	(42.1)
Qualification Level				
Q1	0.2546	0.3824	0.3363	0.3317
	(55.0)	(41.3)	(44.9)	(24.6)
Q2	0.1665	0.3397	0.1880	0.2540
	(45.4)	(40.9)	(31.9)	(22.7)
Q3	0.1230	0.2606	0.1808	0.2328
	(39.1)	(37.5)	(34.9)	(21.4)
Q4	0.0992	0.2352	0.1149	0.2413
	(31.4)	(35.2)	(24.0)	(29.3)
Q5	0.0490	0.1286	0.0450	0.1122
	(17.3)	(26.7)	(11.0)	(23.1)
Q6	0.0416	0.0902	0.0246	0.0702
	(14.0)	(16.9)	(5.8)	(13.6)
Q7	0.0308	0.0366	-0.0009*	0.0547
	(9.3)	(6.3)	(-0.2)	(8.2)
$\overline{R}^2$	0.88	0.88	0.93	0.88
Sum of squared residuals	9601.5	2939.4	2677.2	1511.7
Number of Observations	471408	170414	155112	101750

Notes: (i) t-statistics are in parentheses;
(ii) all estimates are significant at 1%, except those with an *.

5.2.2 Workers' Accessions and Separations

In this Section we examine real wage cyclicality for workers' hiring and separations. Previous research on real wage cyclicality has been investigating the heterogeneity in real wage behavior for workers with a different mobility status, namely job changers and job stayers. However, the issue of heterogeneity for workers' accessions and separations has not yet been explored in the literature. One of the purposes of this study is to investigate how external labor market conditions affect a worker's wage at the moment of hiring and separation. This analysis will enable us to identify the cycle consequences (in terms of wages) for those workes who search for a new job (wage offers) and for those workers who left (voluntary or involuntarily) a job (wage losses). Are real wages sensitive to the conditions prevailing in the labor market when a hiring or separation occurs?

In Tables 1.16 and 1.17 the OLS estimates for newly hired workers are presented separately for both men and women.[27] As mentioned in Section 4.2, the design of the sample is cross-sectiona, making the use of panel data estimators unfeasible.

The overall results indicate a strongly procyclical behavior of real wages of newly hired workers, regardless of considering the sample of men or women or the wage measure being used. According to specification 2 of Table 1.16, an unemployment rate coefficient of -2.08 was obtained for men and a slightly less procyclical one of -1.78 for women. These values are considerably higher than those obtained for continuously employed workers, in particular, for job stayers.

The estimates of Table 1.16 also show the importance of qualifications in wage determination for this sample of workers when compared to continuously employed workers. The effects of education and age on wages are not as strong as for continuously employed workers.

These results suggest that in booms newly hired workers gain access to jobs that pay higher wages. Some alternative explanations have been advanced in order to explain why new hires have more procyclical wages. The more frequent relies on the existence of interindustry wage differentials. This interpretation was first advanced by Okun (1973), who argued that certain jobs offer rents to workers. If these sectors are also more cyclically sensitive, workers can switch into high-paying jobs during booms because such jobs are less tightly rationed during these times. The problem with this explanation is that many of the workers who change jobs in booms leave these jobs in subsequent recessions. Moreover, Okun did not explain what the sources of interindustry wage differentials are.

Recently, Barlevy (2001) offered a new explanation for the existence of more procyclical wages of job changers: compensating differentials. In order to show that compensating differentials instead of interindustry wage differentials generate a more procyclical behavior of wages of changers, Barlevy developed a model that relates unemployment insurance and wage cyclicality. His empirical

[27] In Tables A.1.7 (for men) and A.1.8 (for women) of Appendix A the descriptive statistics of all variables are provided for this dataset.

finding of a negative relationship between wage cyclicality among job changers and the level of unemployment insurance benefits, supports the view that job changers' wages are more procyclical because in booms they obtained jobs that pay a compensating differential for the risk of layoff. In this case, workers who change jobs during booms may not realize true gains from the higher wages they receive, since these gains are typically offset during recessions.

Beaudry and DiNardo (1991) advanced a more convincing explanation for the differences in wage cyclicality between job stayers and job changers, even though their explanation abstracts from heterogeneity across jobs. According to their findings, current unemployment rate does not affect wages after controlling for the best labor market conditions, since a worker was hired at his/her current job. Indeed, when workers are not mobile between employers, current labor market conditions do not affect current wages. In this case, current wages are negatively correlated with the unemployment rate at the time each worker was hired. However, if workers are very mobile, wages are correlated with the best labor market conditions observed since the worker was hired.

Arozamena and Centeno (2001) present a model that combines job matching and business cycle effects and argue that as the employment relationship progresses and specific human capital is accumulated, the worker, by receiving a fraction of the return to that human capital, is gradually insulated from cyclical variations in the external labor market. They found robust empirical evidence supporting this prediction: the elasticity of wages to the unemployment rate decreases with tenure.

The results obtained for the sample of newly hired workers (accessions) seem to be more consistent with these last two explanations. Newly hired workers have more procyclical wages because they hold more unstable job matches with no access to insurance, being subject to the aggregate conditions that prevail in the labor market at the time they start a new job. It is noteworthy that in Portugal, as shown by Varejão and Portugal (2001), accessions and separations occur predominantly in the group of workers with temporary contracts. According to their findings, between 1991 and 1998, fixed-term contracts account for 62% of all accessions and 43% of all separations.

Table 1.16: OLS Results
Accessions
(Dependent variable: log Average Hourly Earnings)

	Men		Women	
Independent Variables	(1)	(2)	(1)	(2)
U	-2.15	-2.08	-1.72	-1.78
	(-13.3)	(-16.1)	(-9.6)	(-12.6)
T	0.0220	0.0171	0.0186	0.0133
	(10.2)	(9.9)	(7.8)	(7.1)
TSQ	0.0010	0.0005	0.0012	0.0006
	(5.9)	(3.5)	(6.4)	(4.2)
AGE		0.0293		0.0217
		(45.8)		(28.3)
AGESQ		-0.0003		-0.0002
		(-36.9)		(-21.7)
EDUC		0.0385		0.0431
		(98.8)		(102.5)
Qualification Level				
Q1		0.8542		0.7876
		(104.1)		(70.9)
Q2		0.7206		0.7184
		(85.8)		(69.9)
Q3		0.5511		0.4752
		(78.2)		(40.8)
Q4		0.5222		0.5577
		(75.9)		(79.9)
Q5		0.2527		0.1731
		(74.0)		(53.4)
Q6		0.1792		0.1277
		(44.5)		(36.3)
Q7		0.1238		0.1372
		(33.7)		(36.7)
CONST	5.130	4.160	4.942	4.122
	(381.9)	(274.6)	(331.5)	(305.0)
$\overline{R}^2$	0.10	0.42	0.10	0.44
Sum of squared residuals	19897.7	12790.7	10529.1	6536.3
Number of Observations	115009		72662	

Notes: (i) t-statistics are in parentheses;
(ii) all estimates are significant at 1%.

Table 1.17: OLS Results (specification 2)
Accessions
Alternative Wage Measures

	Wage Measure			
	AHE	AHEIOT	AHBW	AMBW
Men				
Cycle Regressor (U)	-2.08	-2.08	-2.11	-2.14
	(-16.1)	(-16.1)	(-19.2)	(-15.6)
Women				
Cycle Regressor (U)	-1.78	-1.78	-1.79	-1.46
	(-12.6)	(-12.6)	(-14.1)	(-9.0)

Notes: (i) t-statistics are in parentheses;
(ii) all estimates are significant at 1%.

Tables 1.18 and 1.19 report the same regression results for the sample of recently separated workers (men and women).[28] Once again, the estimates of the unemployment rate coefficient exhibit a procyclical behavior of average hourly earnings. A 1-percentage point increase in the unemployment rate decreases average hourly earnings by 0.6% for men and by about 1.0% for women. Two facts emerge from these results. The first is the existence of a more pronounced cyclical behavior of wages for female separations than for continuously employed females. This figure of 1.0% can be compared to that of 0.45% for women employed for two consecutive years. The second is, in contrast with previous results, the existence of a greater cyclical wage sensitivity for female separations than for male separations.

Both pieces of evidence are likely to be produced by the heterogeneity in the composition of the pool of separations of men and women. The pool of separations may include workers with a low tenure as, for example, those with a fixed-term contract, and may include workers with a higher tenure as, for instance, workers close to retirement. In other situations, separations may be produced by the displacement of workers due to collective dismissals or the shutdown of the plant. In such cases, wage concession mechanisms may be at work.

Low-tenure workers, mostly women, account for a disproportionate share of worker turnover. In fact, in the sample a half of the total separations is produced by workers with three or fewer years of tenure. If we exclude from the separation sample those workers with more than three years of tenure we obtain a higher unemployment regression coefficient (1.2% for both men and women). Or in another way, 33 percent of the separations are also classified as accessions over the one year interval. If we excise these workers from the sample we would obtain much lower unemployment regression coefficients (0.42% for men and 0.69% for women).

[28] In Tables A.1.9 (for men) and A.1.10 (for women) of Appendix A the descriptive statistics of all variables are provided for this dataset.

Table 1.18: OLS Results
Separations
(Dependent variable: log Average Hourly Earnings)

	Men		Women	
Independent Variables	(1)	(2)	(1)	(2)
U	-0.37*	-0.60	-0.72*	-1.02
	(-1.4)	(-2.8)	(-2.5)	(-4.4)
T	0.0517	0.0409	0.0259	0.0205
	(12.5)	(12.8)	(5.8)	(5.8)
TSQ	-0.0018	-0.0018	0.0002*	-0.0003*
	(-4.9)	(-6.4)	(0.4)	(-1.0)
AGE		0.0316		0.0228
		(42.6)		(26.4)
AGESQ		-0.0003		-0.0002
		(-29.3)		(-16.4)
EDUC		0.0471		0.0493
		(91.2)		(89.0)
Qualification Level				
Q1		0.8334		0.7763
		(80.5)		(53.1)
Q2		0.7800		0.7625
		(76.2)		(60.0)
Q3		0.5335		0.5044
		(69.0)		(43.7)
Q4		0.5193		0.5866
		(63.2)		(67.8)
Q5		0.2443		0.1855
		(47.7)		(38.3)
Q6		0.1858		0.1392
		(32.4)		(27.4)
Q7		0.0979		0.1141
		(17.3)		(20.0)
CONST	5.093	3.919	4.969	4.036
	(218.7)	(176.6)	(196.2)	(162.6)
$\overline{R}^2$	0.06	0.44	0.05	0.42
Sum of squared residuals	20952.5	12574.8	10469.5	6388.9
Number of Observations	88894		56979	

Notes: (i) t-statistics are in parentheses;
(ii) all estimates are significant at 1%, except those with an *.

Table 1.19: OLS Results (specification 2)
Separations
Alternative Wage Measures

	Wage Measure			
	AHE	AHEIOT	AHBW	AMBW
Men				
Cycle Regressor (U)	-0.60	-0.58	-0.53	-0.38*
	(-2.8)	(-2.7)	(-2.9)	(-1.8)
Women				
Cycle Regressor (U)	-1.02	-0.96	-0.98	-0.81
	(-4.4)	(-4.1)	(-4.7)	(-3.1)

Notes: (i) t-statistics are in parentheses;
(ii) all estimates are significant at 1%, except those with an *.

6 Conclusion

The aim of this study is to provide further evidence on real wage cyclicality using Portuguese data for the period 1986-98, addressing the issue of heterogeneity in wages responses to aggregate labor market conditions for workers' hirings and separations.

The empirical evidence gathered in this exercise is fourfold. First, there is an indication of a moderate procyclical behavior of real wages for continuously employed workers, in particular, for job stayers. This pattern holds for different measures of wages used and specifications. Second, the comparison of the OLS with the fixed-effects regression results suggests that unobserved heterogeneity plays a non-trivial role in the cyclicality of wages in a way that magnifies the cyclicality of real wages.

Third, the examination of real wage cyclicality of workers' accessions revealed a strong procyclical behavior of wages for this type of worker. This empirical evidence is in accordance with the idea that entry wages are much more procyclical than current wages, which confirms previous empirical research consistent with the hypothesis of implicit shielding agreements between employers and workers in wage determination [Beaudry and DiNardo (1991) and Arozamena and Centeno (2001)]. Four, whereas in the case of accessions the evidence of cyclicality is very strong, for the case of separations a number of conflicting forces seem to be at work. Here it is worth emphasizing that a non-trivial portion of separations are made-up of recent hirings. This particular composition of the separation pool tends to amplify the cyclicality of the wages of separated workers.

In sum, the empirical evidence suggests that even micro-data estimates of real wage cyclicality may conceal a strong procyclical wage behavior, when heterogeneity on wages responses to aggregate conditions between employed workers and hirings and separations are not taken into account.

APPENDIX A - Descriptive Statistics

Table A.1.1: Descriptive Statistics (1986-98)
Long-term Employee/Men

	MEAN	STDV	MIN	MAX
Variables				
Unemployment Rate (U)	6.34	1.45	4.1	8.5
Trend (T)	6.17	3.85	0.0	12.0
Age (AGE)	42.00	8.08	18.0	64.0
Education (EDUC)	5.59	2.73	0.0	16.0
Qualification Level				
Manager and Highly Professional (Q1)	0.015		0.0	1.0
Professional (Q2)	0.028		0.0	1.0
Supervisors (Q3)	0.097		0.0	1.0
Highly Skilled (Q4)	0.066		0.0	1.0
Skilled (Q5)	0.572		0.0	1.0
Semi-skilled (Q6)	0.171		0.0	1.0
Unskilled (Q7)	0.042		0.0	1.0
Apprentices (Q8)	0.008		0.0	1.0
Average Hourly Earnings (AHE)	5.669	0.43	4.35	7.18
Number of Observations	471408			
Average Hourly Earnings inc OT (AHEIOT)	5.681	0.44	4.35	7.18
Avearge Hourly Base Wages (AHBW)	5.454	0.37	4.32	6.84
Average Monthly Base Wages (AMBW)	10.606	0.33	8.99	11.74

Table A.1.2: Descriptive Statistics (1986-98)
Long-term Employee/Women

	MEAN	STDV	MIN	MAX
Variables				
Unemployment Rate (U)	6.34	1.45	4.1	8.5
Trend (T)	6.17	3.85	0.0	12.0
Age (AGE)	39.34	7.86	18.0	64.0
Education (EDUC)	6.20	3.04	0.0	16.0
Qualification Level				
Manager and Highly Professional (Q1)	0.010		0.0	1.0
Professional (Q2)	0.022		0.0	1.0
Supervisors (Q3)	0.045		0.0	1.0
Highly Skilled (Q4)	0.073		0.0	1.0
Skilled (Q5)	0.438		0.0	1.0
Semi-skilled (Q6)	0.321		0.0	1.0
Unskilled (Q7)	0.081		0.0	1.0
Apprentices (Q8)	0.011		0.0	1.0
Average Hourly Earnings (AHE)	5.431	0.50	4.35	7.17
Number of Observations	155112			
Average Hourly Earnings inc OT (AHEIOT)	5.434	0.50	4.35	7.18
Average Hourly Base Wages (AHBW)	5.282	0.45	4.34	6.82
Averahe Monthly Base Wages (AMBW)	10.404	0.40	8.99	11.71

Table A.1.3: Descriptive Statistics (1986-98)
Long-term Employee/Men Job Stayers

	MEAN	STDV	MIN	MAX
Variables				
Unemployment Rate (U)	6.34	1.45	4.1	8.5
Trend (T)	6.17	3.85	0.0	12.0
Age (AGE)	42.59	8.15	18.0	64.0
Education (EDUC)	5.41	2.57	0.0	16.0
Qualification Level				
Manager and Highly Professional (Q1)	0.011		0.0	1.0
Professional (Q2)	0.030		0.0	1.0
Supervisors (Q3)	0.106		0.0	1.0
Highly Skilled (Q4)	0.065		0.0	1.0
Skilled (Q5)	0.573		0.0	1.0
Semi-skilled (Q6)	0.162		0.0	1.0
Unskilled (Q7)	0.046		0.0	1.0
Apprentices (Q8)	0.007		0.0	1.0
Average Hourly Earnings (AHE)	5.630	0.42	4.35	7.18
Number of Observations	285708			
Average Hourly Earnings inc OT (AHEIOT)	5.643	0.43	4.35	7.18
Average Hourly Base Wages (AHBW)	5.423	0.37	4.32	6.84
Average Monthly Base Wages (AMBW)	10.587	0.33	8.99	11.74

Table A.1.4: Descriptive Statistics (1986-98)
Long-term Employee/Women Job Stayers

	MEAN	STDV	MIN	MAX
Variables				
Unemployment Rate (U)	6.34	1.45	4.1	8.5
Trend (T)	6.17	3.85	0.0	12.0
Age (AGE)	39.68	7.96	18.0	64.0
Education (EDUC)	6.11	3.01	0.0	16.0
Qualification Level				
Manager and Highly Professional (Q1)	0.009		0.0	1.0
Professional (Q2)	0.024		0.0	1.0
Supervisors (Q3)	0.042		0.0	1.0
Highly Skilled (Q4)	0.068		0.0	1.0
Skilled (Q5)	0.415		0.0	1.0
Semi-skilled (Q6)	0.341		0.0	1.0
Unskilled (Q7)	0.091		0.0	1.0
Apprentices (Q8)	0.010		0.0	1.0
Average Hourly Base Earnings (AHE)	5.421	0.50	4.36	7.16
Number of Observations	103476			
Average Hourly Earnings inc OT (AHEIOT)	5.424	0.50	4.36	7.18
Average Hourly Base Wages (AHBW)	5.273	0.45	4.35	6.82
Average Monthly Base Wages (AMBW)	10.400	0.40	8.99	11.71

Table A.1.5: Descriptive Statistics (1986-98)
Two-year Employee/Men

	MEAN	STDV	MIN	MAX
Variables				
Unemployment Rate (U)	6.39	1.45	4.1	8.5
Trend (T)	6.39	3.75	0.0	12.0
Age (AGE)	38.0	11.14	18.0	64.0
Education (EDUC)	5.99	3.06	0.0	16.0
Qualification Level				
Manager and Highly Professional (Q1)	0.030		0.0	1.0
Professional (Q2)	0.029		0.0	1.0
Supervisors (Q3)	0.066		0.0	1.0
Highly Skilled (Q4)	0.061		0.0	1.0
Skilled (Q5)	0.529		0.0	1.0
Semi-skilled (Q6)	0.151		0.0	1.0
Unskilled (Q7)	0.089		0.0	1.0
Apprentices (Q8)	0.045		0.0	1.0
Average Hourly Earnings (AHE)	5.483	0.52	4.10	7.46
Number of Observations	170414			
Average Hourly Earnings inc OT (AHEIOT)	5.490	0.53	4.10	7.94
Avearge Hourly Base Wages (AHBW)	5.345	0.46	2.60	7.45
Average Monthly Base Wages (AMBW)	10.489	0.47	6.11	12.65

Table A.1.6: Descriptive Statistics (1986-98)
Two-year Employee/Women

	MEAN	STDV	MIN	MAX
Variables				
Unemployment Rate (U)	6.39	1.41	4.1	8.5
Trend (T)	6.81	3.69	0.0	12.0
Age (AGE)	34.45	10.10	18.0	64.0
Education (EDUC)	6.33	3.11	0.0	16.0
Qualification Level				
Manager and Highly Professional (Q1)	0.014		0.0	1.0
Professional (Q2)	0.017		0.0	1.0
Supervisors (Q3)	0.023		0.0	1.0
Highly Skilled (Q4)	0.050		0.0	1.0
Skilled (Q5)	0.415		0.0	1.0
Semi-skilled (Q6)	0.296		0.0	1.0
Unskilled (Q7)	0.116		0.0	1.0
Apprentices (Q8)	0.070		0.0	1.0
Average Hourly Earnings (AHE)	5.220	0.48	4.12	7.44
Number of Observations	101750			
Average Hourly Earnings inc OT (AHEIOT)	5.222	0.48	4.12	7.48
Avearge Hourly Base Wages (AHBW)	5.128	0.44	3.09	7.42
Average Monthly Base Wages (AMBW)	10.235	0.47	6.01	12.64

Table A.1.7: Descriptive Statistics (1986-98)
Accessions/Men

	MEAN	STDV	MIN	MAX
Variables				
Unemployment Rate (U)	6.40	1.44	4.1	8.5
Trend (T)	6.13	3.91	0.0	12.0
Age (AGE)	30.91	10.08	18.0	64.0
Education (EDUC)	6.11	3.00	0.0	16.0
Qualification Level				
Manager and Highly Professional (Q1)	0.020		0.0	1.0
Professional (Q2)	0.017		0.0	1.0
Supervisors (Q3)	0.025		0.0	1.0
Highly Skilled (Q4)	0.027		0.0	1.0
Skilled (Q5)	0.460		0.0	1.0
Semi-skilled (Q6)	0.128		0.0	1.0
Unskilled (Q7)	0.195		0.0	1.0
Apprentices (Q8)	0.128		0.0	1.0
Average Hourly Earnings (AHE)	5.188	0.44	3.90	7.36
Number of Observations	115009			
Average Hourly Earnings inc OT (AHEIOT)	5.192	0.44	3.90	7.37
Avearge Hourly Base Wages (AHBW)	5.092	0.39	3.87	7.23
Average Monthly Base Wages (AMBW)	10.205	0.44	7.95	12.29

Table A.1.8: Descriptive Statistics (1986-98)
Accessions/Women

	MEAN	STDV	MIN	MAX
Variables				
Unemployment Rate (U)	6.35	1.39	4.1	8.5
Trend (T)	6.72	3.84	0.0	12.0
Age (AGE)	29.34	9.11	18.0	64.0
Education (EDUC)	6.71	3.17	0.0	16.0
Qualification Level				
Manager and Highly Professional (Q1)	0.012		0.0	1.0
Professional (Q2)	0.014		0.0	1.0
Supervisors (Q3)	0.010		0.0	1.0
Highly Skilled (Q4)	0.032		0.0	1.0
Skilled (Q5)	0.309		0.0	1.0
Semi-skilled (Q6)	0.212		0.0	1.0
Unskilled (Q7)	0.188		0.0	1.0
Apprentices (Q8)	0.224		0.0	1.0
Average Hourly Earnings (AHE)	5.036	0.40	3.88	7.37
Number of Observations	72662			
Average Hourly Earnings inc OT (AHEIOT)	5.038	0.40	3.88	7.37
Average Hourly Base Wages (AHBW)	4.970	0.37	3.88	7.22
Average Monthly Base Wages (AMBW)	10.037	0.42	7.92	12.30

Table A.1.9: Descriptive Statistics (1986-97)
Separations/Men

	MEAN	STDV	MIN	MAX
Variables				
Unemployment Rate (U)	6.45	1.56	4.1	8.5
Trend (T)	5.88	3.72	0.0	11.0
Age (AGE)	35.38	12.14	18.0	64.0
Education (EDUC)	5.73	2.88	0.0	16.0
Qualification Level				
Manager and Highly Professional (Q1)	0.024		0.0	1.0
Professional (Q2)	0.022		0.0	1.0
Supervisors (Q3)	0.049		0.0	1.0
Highly Skilled (Q4)	0.039		0.0	1.0
Skilled (Q5)	0.497		0.0	1.0
Semi-skilled (Q6)	0.141		0.0	1.0
Unskilled (Q7)	0.138		0.0	1.0
Apprentices (Q8)	0.090		0.0	1.0
Average Hourly Earnings (AHE)	5.287	0.50	3.88	7.52
Number of Observations	88894			
Average Hourly Earnings inc OT (AHEIOT)	5.290	0.50	3.88	7.52
Average Hourly Base Wages (AHBW)	5.183	0.45	3.87	7.41
Average Monthly Base Wages (AMBW)	10.299	0.48	7.93	12.43

Table A.1.10: Descriptive Statistics (1986-97)
Separations/Women

	MEAN	STDV	MIN	MAX
Variables				
Unemployment Rate (U)	6.37	1.52	4.1	8.5
Trend (T)	6.32	3.62	0.0	11.0
Age (AGE)	33.09	10.65	18.0	64.0
Education (EDUC)	6.17	3.05	0.0	16.0
Qualification Level				
Manager and Highly Professional (Q1)	0.011		0.0	1.0
Professional (Q2)	0.015		0.0	1.0
Supervisors (Q3)	0.018		0.0	1.0
Highly Skilled (Q4)	0.040		0.0	1.0
Skilled (Q5)	0.368		0.0	1.0
Semi-skilled (Q6)	0.271		0.0	1.0
Unskilled (Q7)	0.148		0.0	1.0
Apprentices (Q8)	0.130		0.0	1.0
Average Hourly Earnings (AHE)	5.097	0.44	3.89	7.49
Number of Observations	56979			
Average Hourly Earnings inc OT (AHEIOT)	5.099	0.44	3.89	7.49
Average Hourly Base Wages (AHBW)	5.026	0.41	3.87	7.40
Average Monthly Base Wages (AMBW)	10.103	0.46	7.91	12.40

References

[1] Abraham, K. and J. Kaltiwanger (1995), "Real Wages and the Business Cycle", *Journal of Economic Literature*, Vol. 33, pp. 1215-1264.

[2] Arozamena, L. and M. Centeno (2001), "Tenure, Business Cycle and Wage-Setting Process", Banco de Portugal Working Paper 8-01.

[3] Banco de Portugal (1999), *Relatório Anual de 1998*, Lisboa: Banco de Portugal.

[4] Barlevy, G. (2001), "Why Are the Wages of Job Changers So Procyclical?", *Journal of Labor Economics*, Vol. 19, pp. 837-878.

[5] Barro, R. and H. Grossman (1971), "A General Disequilibrium Model of Income and Employment", *American Economic Review*, Vol. 61, pp. 82-93.

[6] Beaudry, P. and J. DiNardo (1991), "The Effect of Implicit Contracts on the Movement of Wages over the Business Cycle: Evidence from Micro Data", *Journal of Political Economy*, Vol. 99, pp. 665-688.

[7] Bils, M. (1985), "Real Wages over the Business Cycle: Evidence from Panel Data", *Journal of Political Economy*, Vol. 93, pp. 666-689.

[8] Blank, R. (1990), "Why are Wages Cyclical in the 1970s?", *Journal of Labor Economics*, Vol. 8, pp. 16-47.

[9] Blundell, R., H. Reed and T. Stoker (2000), "Interpreting Aggregate Wage Growth", IFS Publication.

[10] Bodkin, R. (1969), "Real Wages and Cyclical Variations in Employment: a Re-examination of the Evidence", *Canadian Journal of Economics*, Vol. 2, pp. 353-374.

[11] Brandolini, A. (1995), "In Search of a Stylised Fact: Do Real Wages Exhibit a Consistent Pattern of Cyclical Variability?", *Journal of Economic Surveys*, Vol. 9, pp. 103-163.

[12] Burda, M. (1985), "New Evidence on Real-Wage Employment Correlation from U.S. Manufacturing Data", *Economics Letters*, Vol. 18, pp. 283-285. Errata: *Economics Letters*, Vol. 19, 1985, p. 399.

[13] Canzoneri, M. (1978), "The Returns of Labor and the Cyclical Behavior of Real Wages", *Review of Economics and Statistics*, Vol. 60, pp. 19-24.

[14] Chirinko, R. (1980), "The Real Wage Rate over the Business Cycle", *Review of Economics and Statistics*, Vol. 62, pp. 459-461.

[15] Devereux, P. (2001), "The Cyclicality of Real Wages within Employer-Employee Matches", *Industrial and Labor Relations Review*, Vol. 54, pp. 835-850.

[16] Dunlop, J. (1938), "The Measurement of Real Wages and Money Wage Rates", *Economic Journal*, Vol. 48, pp. 413-434.

[17] Geary, T. and J. Kennan (1982), "The Employment-Real Wage Relationship", *Journal of Political Economy*, Vol. 90, pp. 854-871.

[18] Grant, D. (2001), "A Comparison of the Cyclical Behavior of Union and Nonunion Wages in the United States", *Journal of Human Resources*, Vol. 36, pp. 31-57.

[19] Keane, M., R. Moffitt and D. Runkle (1988), "Real Wages over the Business Cycle: Estimating the Impact of Heterogeneity with Micro Data", *Journal of Political Economy*, Vol. 96, pp. 1236-1266.

[20] Keynes, J. (1936), *The General Theory of Employment, Interest and Money*, London: Macmillan.

[21] Keynes, J. (1939), "Relative Movements of Real Wages and Output", *Economic Journal*, Vol. 49, pp. 34-51.

[22] Meghir, C. and E. Whitehouse (1996), "The Evolution of Wages in the United Kingdom: Evidence from Micro Data", *Journal of Labor Economics*, Vol. 14, pp. 1-25.

[23] Mehra, Y. (1982), "Real Wages and Employment: Evidence from Disaggregated Data.", *Eastern Economic Journal*, Vol. 8, pp. 191-196.

[24] Michie, J. (1987), *Wages in the Business Cycle*, London: Printer.

[25] Mitchell, M., M. Wallace and J. Warner (1985), "Real Wages over the Business Cycle: Some Further Evidence", *Southern Economic Journal*, Vol. 51, pp. 1162-1173.

[26] Neftci, S. (1978), "A Time-Series Analysis of the Real Wages-Employment Relationship", *Journal of Political Economy*, Vol. 86, pp. 281-291.

[27] Okun (1973), "Upward Mobility in a High-pressure Economy", *Brookings Papers on Economic Activity*, Vol. 1, pp. 207-252.

[28] Otani, I. (1978), "Real Wages and Business Cycles Revisited", *Review of Economics and Statistics*, Vol. 60, pp. 301-304.

[29] Raisian, J. (1983), "Contracts, Job Experience, and Cyclical Labor Market Adjustments", *Journal of Labor Economics*, Vol. 1, pp. 152-170.

[30] Rosen, S. (1985), "Implicit Contracts: a Survey", *Journal of Economic Literature*, Vol. 23, pp. 1144-1175.

[31] Sargent, J. (1978), "Estimation of Dynamic Labor Demand Schedules Under Rational Expectations", *Journal of Political Economy*, Vol. 86, pp. 1009-1044.

[32] Shin, D. (1994), "Cyclicality of Real Wages among Young Men", *Economics Letters*, Vol. 46, pp. 137-142.

[33] Solon, G., R. Barsky and J. Parker (1994), "Measuring the Cyclicality of Real Wages: How Important is Composition Bias?", *Quarterly Journal of Economics*, Vol. 109, pp. 1-25.

[34] Sumner, S. and S. Silver (1989), "Real Wages, Employment, and the Phillips Curve", *Journal of Political Economy*, Vol. 97, pp. 706-720.

[35] Tarshis, L. (1939), "Changes in Real and Money Wages", *Economic Journal*, Vol. 49, pp. 150-154.

[36] Tatom, J. (1980), "The 'Problem' of Procyclical Real Wages and Productivity", *Economic Journal of Political Economy*, Vol. 88, pp. 385-394.

[37] Tremblay, C. (1990), "Wage Patterns of Women over the Business Cycle", *Quarterly Review of Economics and Business*, Vol. 30, pp. 90-101.

[38] Varejão, J. and P. Portugal (2001), "Why Do Firms Use Fixed-Term Contracts?", *mimeo.*

[39] Ziliak, J., B. Wilson and J. Stone (1999), "Spatial Dynamics and Heterogeneity in the Cyclicality of Real Wages", *Review of Economics and Statistics*, Vol. 81, pp. 227-236.

Chapter 2

Insider Power and Wage Determination: Evidence from a Panel of Large Firms

1 Introduction

Since the early 1970s Western Europe has suffered a period of high and rising unemployment. In 1975, the unemployment rate in the European Union averaged close to 4 percent. In 1985, it had increased to 10 percent. In 2000, it stands at 8.2 percent, which is about three percentage points below its peak of 11.1% in 1994.[1] Despite high and rising unemployment, real wages have risen sharply in Europe since the mid-1970s. Moreover, since that time the broad pattern of events is one in which inflation and unemployment have moved together.

Traditionally, macroeconomics has been concerned with how temporary shocks make unemployment fluctuate in the short term around its equilibrium level (the NAIRU), while labor economics has focused on what determines that equilibrium level. But it has become more and more obvious that the equilibrium level itself varies greatly between decades, with previous unemployment exerting a persistent effect on subsequent unemployment. To explain this persistence, new microeconomic foundations of macroeconomics were required, going far beyond the influences considered in the 1970s. In fact, standard theories of aggregate demand and supply do not easily explain such persistent effects on the level of unemployment. Theories that allow for hysteresis - a strong dependence of current unemployment on past employment - seem to be the more appropriate ones.

Two types of explanations were advanced in order to explain the high and persistent unemployment in Europe in the last two decades. The first explanation emphasizes the role of human capital accumulation as explored, among others, by Phelps (1972), Bruno and Sachs (1985) and Pissarides (1992). The human capital argument holds that workers who are unemployed lose the opportunity to maintain and update their skills by working. Prolonged unemployment depreciates human capital skills, and it becomes more difficult to find a new job. However, direct and indirect evidence does not support the idea that skill deterioration associated with unemployment is strong enough to explain why the long-term unemployed remain unemployable [Blanchard (1991)]. A new version of this story rests on the attitudes of employers toward the long-term unemployed [see Blanchard (1991) and Blanchard and Diamond (1994)]. The idea

[1] These figures are from Eurostat and the average unemployment rate refers to the countries that are members of the European Union.

is that firms faced with a queue of job applicants of different unemployment duration, may hire first those who have become unemployed more recently. The consequences of this behavior is that high rates of unemployment may not greatly concern the currently employed workers, who realize that their chances of finding employment, were they to become unemployed, are much higher than those of the average unemployed worker. In this context of ranking, and as long as employed workers realize that their reemployment probabilities differ more and more from those who remain unemployed, a high unemployment level may not exert substantial downward pressure on wages, leading to unemployment persistence.

Finally, unemployment persistence may arise when outsiders' effort into job search falls with their duration of unemployment. As the duration of unemployment lengthens, more and more workers stop searching, leading to their exclusion from the labor market and becoming irrelevant to the process of wage determination. Once again, current unemployment gives rise to future unemployment.

The second explanation emphasizes not the characteristics of the unemployed, but the behavior of the insiders. The insider-outsider explanation is based on the idea that the level of wages is primarily determined by the currently employed workers (the so-called 'insiders'), with unemployed (the 'outsiders') playing little or no role in the process of wage bargaining. Blanchard and Summers (1986) and Gottfries and Horn (1987) were the first to analyze the implications of an insider-outsider approach for time dependence of unemployment (i. e, dependence of current unemployment on past unemployment). The basic idea is that different membership effects between insiders who are employed and outsiders who want jobs may generate unemployment persistence. Wages are set by bargaining between firms and currently employed workers in order to guarantee that insiders have a high probability of remaining employed in the face of random demand shocks. In fact, insiders are concerned about maintaining their jobs, not ensuring the employment of outsiders. This has two main implications, as observed by Blanchard and Summers (1986). First, in the absence of shocks, any level of employment of insiders is self-sustaining. Second, in the presence of shocks, employment follows a process similar to a random walk. If, for example, a negative demand shock occurs in the previous period, some insider workers will lose their jobs. Then, since there are fewer workers employed, wages can be raised in the current period without any substantial increase in the likelihood that an employed worker loses his job. Since higher wages imply that fewer unemployed workers will be hired, an adverse shock may have a persistent negative effect on (un)employment.

In order to provide empirical evidence on hysteresis theories, Blanchard and Summers (1986) estimated aggregated wage equations for the United Kingdom, France, Germany, and the United States for the period 1953-84. They found that the United States, contrary to Europe, exhibited much less hysteresis. Both unemployment and employment are more persistent in Europe than in the United States. The main issue, however, is to determine the factors that cause hysteresis in Europe. It is yet unclear whether the cause of hysteresis in Europe

is the result of institutional factors, in particular the wage setting system, or the result of the sequence of adverse shocks that has caused high unemployment.

Relying on the distinction between insider and outsider workers, the insider-outsider theories of wage formation aimed to explain why wages may be set above their market-clearing levels.[2] Furthermore, this approach attempts to explain why unemployed workers do not compete for existing jobs by offering to work at jobs for which they are qualified at a wage lower than that currently being paid to incumbents. Lindbeck and Snower (1986) showed that the existence of costs associated with insider-outsider turnover might explain why firms do not replace their high-wage insiders with low-wage outsiders. Accordingly, involuntary unemployment can arise due to the existence of turnover costs such as hiring, training and firing costs or the costs generated by the disincentive to cooperate with outsiders, that make it costly to the firm to replace an insider worker with an unemployed worker.[3] The rent associated with these turnover costs gives some labor market power to insiders in the process of wage negotiation.

In sum, under insider wage setting theories a firm's wage should rise when an autonomous increase in its productivity (or in the demand for its product or in its output price) occurs. Under these circumstances the incentives for the firm to expand output and employment are much reduced. As stated by Solow (1985): "...one reason for the persistence of unemployment over a wide range of fluctuations of aggregate demand is the willingness and ability of insiders to convert higher demand into higher wages for themselves rather than into increased access to jobs for outsiders [Solow (1985), p. 427].

The evidence of insider power led to the belief that wages might be widely influenced by firm's internal conditions rather than by external conditions in the labor market. In fact, the insider-outsider approach relies on the assumption that wages are set through bargaining, not between firms and the whole labor force, but rather between firms and their workers. Gregory (1986) suggests that "...perhaps it is the labor utilization rate within the firm that is particularly important for wage negotiations, rather than the labor utilization rate within the economy." [Gregory (1986), p. S63].

In the 1990s, the microeconometric models of wage determination started including measures of the firm's profits or financial performance as explanatory variables. This literature has focussed directly on rent-sharing models [see, among others, Nickell and Wadhwani (1990), Holmlund and Zetterberg (1991), Abowd and Lemieux (1993), Blanchflower *et al.* (1996) and Hildreth and Oswald (1997)]. These studies used panel data at both firm or industry level and estimated versions of the wage equation with rents per worker included. Albeit using different models of collective bargaining, the results of these studies indicate, in general, that changes in profitability are shown to feed through into long-run changes in wages. A branch of this literature, to which this paper is more closely based, has been focussing on the relative importance

[2] See, for example, Lindbeck and Snower (1985, 1986 and 1988) and Solow (1985).

[3] For a description of this type of costs see Lindbeck and Snower (1986).

of insider versus outsider forces in wage determination [see, for instance, Nickell and Wadhwani (1990), Holmlund and Zetterberg (1991) and Bentolila and Dolado (1994)].[4] They show that firm specific factors, as well as general labor market conditions, have an important role on the process of wage determination, although the weight attached to firm-specific factors varies considerably across countries. Indeed, the results suggest that firm-specific factors are relatively unimportant in the Nordic Countries, of middle importance in Britain, West Germany, Spain and the Netherlands, and highly important in Australia, Canada and the United States.

The main difficulty with this new evidence at the micro level is to show whether the role of firm-specific factors in wage determination is strong enough to generate the aggregate persistence of unemployment that is observed in macro-data.

Using an approach inspired in the one first developed by Nickell and Wadhwani (1990), this study examines the role of insider forces (such as productivity and market share) in contrast with the role of outsider forces (unemployment rate) in the process of wage determination in the Portuguese labor market. For this purpose, we use data from around 900 large Portuguese firms from all sectors over the period 1993-97.

More specifically, this study has two main objectives. First, to evaluate the role and weight of insider forces in wage determination. The Portuguese industrial relations system presents some contrasting features. On one hand, the role of massive wage-setting mechanisms and the existence of extension mechanisms point to a centralized bargaining system. On the other hand, the scattered nature of union organization, the possibility opened to employers to bargain at the firm level, and the presence of a significant wage drift,[5] highlight aspects of decentralization that may grant employers some room of maneuver to set wages.[6]

It is also widely agreed that the weight attached to insider forces in wage determination is higher in decentralized than in centralized economies.[7] According to OECD (1997), Portugal shifted, between 1980 and 1994, from a low to a middle degree of centralization and coordination [see OECD (1997), p. 78]. In this context, it will be interesting to find to what extent the Portuguese system of collective bargaining dominates or not the process of wage formation, leaving no scope to negotiations at the firm level.

The second objective is to test empirically to what extent the existence of labor turnover costs creates market power that is exploited by incumbent workers through wage claims. Even though one of the predictions of the insider-outsider theory is that turnover costs make insiders and outsiders imperfect substitutes in the wage bargaining process, the empirical literature on this issue has made

[4] In section 3 an exhaustive description of these studies will be made.

[5] The wage drift measures the difference between the current wage and the wage agreed at the collective bargaining.

[6] See the Annex for a more exhaustive description of the Portuguese Labor Market.

[7] Nickell (1999), for example, calculated a correlation coefficient of -0.92 between the weight of insider effects and the standard Calmfors-Driffill index of centralization.

few attempts to measure the impact of these costs on wage determination. This lack in the insiders-outsiders empirical literature may be explained by the difficulty in measuring directly labor turnover costs (such as severance pay costs, training costs, costs of strikes, costs associated with non-cooperation and harassment activities, and so on). Using a set of proxies that measure insiders' bargaining power, in this study we examine the extent to which labor turnover costs give the insiders market power to extract rents in the form of higher wages.

According to OECD (1999) rankings, Portugal is classified as having a high stricter employment protection legislation. This stricter legislation is associated with lower turnover (high adjustment costs) in the labor market, with both jobs and unemployment spells tending to last longer. As pointed out by Blanchard and Portugal (2001), even though the unemployment rate in Portugal over the past 15 years is quite similar to the U.S. rate, the average unemployment duration in Portugal is more than three times that of the U.S. Symmetrically, flows of workers into unemployment are three times lower in Portugal. In the context of the existence of a stagnant pool of unemployed workers, with low rates of flow into and out of employment, and long unemployment duration, insider-outsider formulations seem to be quite relevant. In fact, a labor market characterized by a high employment protection with high adjustment costs and lower turnover, tends to create market power for currently employed workers that they can exploit to their own advantage. Of course, it can be argued that if turnover costs give bargaining power to workers, and thus the possibility to extract some rents, firms may try to extract those rents from the insiders. However, firms are generally unable to pass these costs on fully to their insiders because firms do not incur these costs until they replace their insiders with new entrants. Moreover, even if firms could extract some of these rents by imposing lump sum payments to insiders upon voluntary quitting or firing "without cause", such fees are usually illegal and incentive-incompatible (Lindbeck and Snower, 2002).

This paper will be organized as follows. The theoretical and empirical frameworks are presented in Section 2. Section 3 presents a brief review of the empirical literature in this area. The data and estimation method are described in Section 4. The empirical results are summarized in Section 5. Section 6 concludes.

2 The Model

2.1 Theoretical Framework

The model used here as a basis for the estimations follows closely those developed in Nickell and Wadhwani (1990), Layard *et al.* (1991), Nickell *et al.* (1994) and Nickell (1999). In these types of models it is presumed that wages are determined through negotiations between the firm and the union, ignoring the intriguing question of the interaction between negotiations at the central, industry, and

local levels. In Portugal negotiations between unions and firms, at the industry level, play an important part in the determination of wages. Hence, modeling the process of wage formation at the firm level using a bargaining approach seems to be an appropriate choice for the Portuguese case.

The model assumes the existence of imperfect competition in the product market. Output demand is subject to exogenous shocks, which are unknown at the moment of the bargaining. The objective of the firm is to maximize its expected profit. It is assumed that the union is only concerned with existing workers and wishes to maximize the expected utility of its median voter. The bargaining is only over wages. Price, output and employment level are set by the firm after the wage negotiations and after the size of the shock is identified.

Under these assumptions, and using the Nash-bargaining *maximand*, the optimal bargain is that which maximizes:

$$\max_{w_i}[(W_i-A)^{\beta}S(W_i)^{\beta}\Pi_i^e(W_i)] \qquad (2.1),$$

where W_i is the real wage at firm i, β is a measure of union power, Π_i^e is the expected profit, A is the expected real income of a worker who loses his job in the firm, and $S(W_i)$ the probability that an employee retains his job in firm i ($S\prime<0$). The bargained wage must satisfy:

$$W_i = \frac{\gamma(W_i)}{\gamma(W_i)-1}A \qquad (2.2),$$

where

$$\gamma(W_i) = \varepsilon_{sw}(W_i)+\frac{W_i N_i^e(W_i)}{\beta_i \pi^e(W_i)} \qquad \text{and} \qquad \varepsilon_{sw} = -\frac{W_i}{S}\frac{\partial S(W_i)}{\partial W_i} \qquad (2.3),$$

N_i^e being the expected level of employment and ε_{sw} the absolute elasticity of the probability of retaining a job with respect to the wage. The expected 'alternative' income, A, is given by

$$A = [1-\varphi(u,z_s)]W^e+\varphi(u,z_s)B = W^e[1-\varphi(u,z_s)(1-b)] \qquad (2.4),$$

where B is the real unemployment benefit, W^e is the expected outside real wage and $b = B/W^e$, is the benefit replacement rate. φ reflects the probability of a displaced worker not obtaining employment elsewhere and is increasing in aggregate unemployment (u). The variable z_s reflects all factors that raise the chances of a displaced worker obtaining alternative employment at given levels of u.

Given equation (2.4) for the expected 'alternative' income, the wage equation (2.2) can be rewritten in the following way:

$$\frac{W_i}{W^e} = \frac{\gamma[1-\varphi(u,z_s)(1-b)]}{\gamma-1} \qquad (2.5).$$

It is assumed that the firm acts as a Cournot industry and produces one good at a real price P_j and has an inverse demand curve

$$P_j = \widetilde{\theta}(\sum_{i\in j} Y_i/Y_{dj})^{\frac{-1}{\eta_j}} \qquad (2.6),$$

where $\widetilde{\theta}$ is a random variable which is revealed after wage bargaining, Y_i is the output of the ith firm, Y_{dj} is an industry demand index. It is also supposed that the production function is Cobb-Douglas:

$$Y_i = \Phi_i N_i^{\alpha} K_i^{1-\alpha} \qquad (2.7),$$

where Φ_i is an index of technical progress and K_i is capital stock, both predetermined. Using (2.6) and (2.7), we obtain the following expressions for γ and ε_{sw}:

$$\gamma = \varepsilon_{sw} + \alpha k_i/(1-\alpha k_i)\beta_i \qquad (2.8),$$

$$\varepsilon_{sw} = f[\frac{W_i}{P_j^e}\frac{1}{\Phi_i \alpha k_i}(\frac{N_{Ii}}{k_i})^{1-\alpha}] \qquad (f'>0) \qquad (2.9).$$

k_i is an indicator of the product market competition defined as:

$$k_i = 1 - MS_i/\eta_j \qquad (2.10).$$

MS_i is market share ($Y_i/\Sigma Y_i$), and N_{Ii} represents the number of employees who participate in wage bargaining, the so-called 'insiders'.

On the basis of (2.5), (2.8), (2.9) and (2.10) and after a logarithmic transformation, we obtain a wage equation of the form:

$$w_i = \lambda[p_j^e + \Phi_i + (1-\alpha)(k_i - n_{Ii})] + (1-\lambda)(w^e - c_1 u + c_2 z_s + c_3 b) + \\ + c_4 MS_i + c_5 \beta_i \qquad (2.11),$$

with $c_1 > 0, c_2 > 0, c_3 > 0$ and $c_5 > 0$.

The interpretation of equation (2.11) is straightforward. The average wage in the ith firm depends on firm-specific factors, outside factors, the firm's market power and workers' bargaining power. The parameter λ may be termed the 'insider weight', and reflects the extent to which wages at the firm level adjust to actual or prospective company performance. In particular, wages depend on

the expected industry output price and technical progress within the firm, as well as on the fixed capital stock relative to the number of employees involved in the wage bargain. λ is increasing in union power (β) and decreasing in the degree of product market competition (k).

The effect of market share on wages is ambiguous. Market power generates monopoly rents and these may be captured by the employees in the form of higher wages. However, as monopoly power increases, this can lead to output and employment restrictions, thereby reducing the probability of job retention and consequently the pressure on wages from the union bargainers. Overall, there are good reasons to believe that survival effects are small, and so, we may expect a positive effect on wages from the market share variable (i. e., $c_4 > 0$), [Nickell, (1999)].

2.2 The Empirical Model

Equation (2.11) constitutes the basis of the empirical analysis. However, some additional transformations are needed in order to estimate the model empirically. First, it will be assumed that the number of insiders is equal to the number of employees in the previous period, i. e., $n_{Ii} = n_{-1i}$. In this case, the first member of equation (2.11) becomes:

$$\begin{aligned} p_j^e + \Phi_i + (1-\alpha)(k_i - n_{Ii}) &= p_j^e + \Phi_i + (1-\alpha)(k_i - n_{-1i}) \\ &= p_j^e + \Phi_i + (1-\alpha)(k_i - n_i) + (1-\alpha)\Delta n_i \\ &= (p_j^e + y_i - n_i) + (1-\alpha)\Delta n_i \end{aligned}$$

The second adjustment implies that the variables b (benefit replacement ratio) and z_s (any factors that reduce competition for jobs among the unemployed) will be omitted from the model. In fact, given that our period of analysis is too short, the figures for these aggregate variables are virtually constant over that period and it does not seem feasible to include them as explanatory variables in the model.

In the third adjustment, and since there are a set of reasons to believe that past wages affect current wages, dynamics will be introduced in the model with the inclusion of the lagged dependent variable as an explanatory variable.

Finally and in fourth place, it is ensured that the long-run homogeneity assumption in the inside and outside factors is verified [$\lambda + (1-\lambda) = 1$]. Equation (2.11) is rearranged by subtracting the aggregate wage rate from both the firm wage rate per employee and the firm value added per employee.

Following these adjustments, and after replacing expected values with the corresponding actual values (innovations are consigned to the error term), equation (2.11) may be rewritten as:

$$w_{it} - \overline{w}_t = a_{0i} + a_1(w_{it-1} - \overline{w}_t) + (1 - a_1)\{\lambda[(p_{jt} + y_{it} - n_{it} - \overline{w}_t) + \\ +(1-\alpha)\Delta n_{it}] + (1-\lambda)c_1 u_t + c_4 MS_{it} + c_5 \beta_{it}\} + \varepsilon_{it} \qquad (2.12).$$

w_{it} is the firm's wage, $(p_{jt}+y_{it}-n_{it})$ is value added per employee, Δn_{it} measures employment change, $\overline{w}_t$ is the outside wage, u_t is the aggregate unemployment, MS_{it} is market share, β_{it} measures union power, ε_{it} is the error term and a_{0i} a firm fixed effect (i=firm, j=industry, t=time).

The interpretation of this equation is also straightforward. Firm's average wage per employee depends on previous wages, on firm-specific factors such as the value-added per employee, the change in the number of insiders and the market share, on outside factors such as the unemployment rate and on workers' bargaining power.[8] λ is the long-run elasticity of firm wages relative to firm revenue per employee.

3 Previous Research

Based on a theoretical framework similar to the one exposed in the previous section, a number of econometric studies emerged in the 1990s investigating the role of insider and outsider forces in wage determination using data at firm or industry level. Internal factors usually considered are the productivity of workers, output prices or profits, union power, market power and lagged wages. The external factors usually include the aggregate unemployment rate and the outside/alternative wage. Measures of unemployment benefits and long-term unemployment are also included as outside factors.

Three main concerns are addressed in this empirical literature. The first is to evaluate the weight of firm-specific factors in the process of wage determination. The empirical literature has been showing that firm-specific factors have an important role in the process of wage determination, although the weight attached to these factors varies significantly across countries. The second is to find evidence in favor of the existence of hysteresis effects. Concerning this issue, the empirical evidence is rather mixed and unsettled. The third is to test the existence of asymmetric effects in wage adjustments. In fact, some evidence has been found that wage adjustments are asymmetric, in the sense that real wages rise more readily in upturns than they fall in downturns.

Next we will present a description of the main results obtained in these empirical studies.

Nickell and Wadhwani (1990) estimated a firm-level equation based on (2.12) using data on a panel of 219 U.K. quoted manufacturing companies for the period 1975-82. They found a significant role in wage determination not only for own price and productivity, but also for the measures of the financial health of the company. Using different instrument sets, different exogeneity assumptions

[8] In order to impose the homogeneity condition, all nominal variables such as w_{it}, w_{it-1} and $(p_{jt} + y_{it} - n_{it})$ are measured as deviations from outside wage.

and alternative samples, they obtained a range of values of the insider weight, λ, between 0.08 and 0.15. No robust evidence was found in favor of insider-based hysteresis effects. In their study, outsider factors, in particular the aggregate unemployment and the proportion of long-term unemployed, also played an important role in wage determination at the firm level.

Nickell and Wadhwani also investigated the possible existence of asymmetric effects in wage responses to demand shocks. The results show some evidence that wages become more rigid when demand is expected to fall.

Nickell *et al.* (1994), using data from 814 British manufacturing companies over the 1975-86 period, extended the results of Nickell and Wadhwani (1990) by focusing on the role of market power in wage determination. They found that product market power has a positive impact on wages. Although the weight of insider forces in wage determination is not influenced by product market power, the role of the external labor market seems to differ with market power. In fact, product market power reduces the generally strong negative effect of unemployment on wages.

Bentolila and Dolado (1994) applied the standard insider-outsider model to the Spanish case by splitting employed individuals into two groups, the permanent employees, which represent the incumbent workforce, and the temporary workforce, who have the characteristics belonging to the outsiders.

The main result of their paper is that a higher share of temporary work leads to an increase in the wages of permanent workers. The sample of 1167 manufacturing firms over the period 1985-88, tends to confirm that the interests of temporary workers are basically disregarded in wage bargains, and that each percentage point of increase in temporary employment could imply up to one-third of one percent estimated increase in wages of permanent employees.

They also present robust evidence of a membership hysteresis effect, i. e., that a reduced union membership lowers the layoff risk for the remaining insiders and thus makes them push harder for wage increases. The novelty in this study is that the membership effect is captured by the change in permanent employment rather than in total employment.

Forslund (1994) and Wulfsberg (1997) investigated the strength of insider effects in wage setting for Sweden and Norway, respectively. These countries are characterized by high degrees of centralization in wage bargaining and, in this context, it should be expected that firm-level performance has a minor influence on wages. In fact, the results obtained for a sample of manufacturing firms revealed a small insider effect for these two countries (the estimates of λ fall in an interval between 0.05 and 0.07).[9]

Concerning hysteresis effects, Forslund found evidence of a positively signed but insignificant membership effect. Conversely, Wulfsberg obtained a significant negative sign for the insider hysteresis effect.

Using industry-level data, Holmlund and Zetterberg (1991) investigated the determinants of industry wages in five countries with very different institutional

[9] Forslund used a sample of 128 manufacturing firms for the years of 1984-88 and Wulfsberg used a sample of 7,323 manufacturing firms for the period 1972-88.

bargaining arrangements: Sweden, Norway, Finland, Germany and the United States. The main results are the following. Industry wages in the U.S. are the most responsive to sectoral price and produtivity changes (this wage response is reinforced by union power). By contrast, industry wages in the Nordic countries are largely unaffected by sectoral conditions. On the contrary, industry wages follow outside wages quite closely in the Nordic countries, and less closely in the U.S. No evidence was found in favor of the hypothesis that wages are inversely related to the size of the incumbent workforce.

Nickell and Kong (1992) also investigated the determinants of industry wages using data on fourteen Bristish production industries for the period of 1961 to 1985. The data revealed that insider forces are important and that insider forces are stronger and outsider forces weaker when unions are powerful. Finally, hysteresis effects arising from insider power are only significant in a reduced number of sectors.

Johansen (1996), using panel data for 117 Norwegian industries from 1966 to 1987 found that industry wages in Norway are significantly affected by industry performance. In fact, he obtained a value of the long-run insider weight for the full sample in the neighborhood of 20%, which is about, respectively, five and three times as high as the corresponding estimate for Norway reported by Holmlund and Zetterberg (1991) and Wulfsberg (1997). As mentioned by the author, the difference between his result and the one reported in Wulfsberg can be justified by aggregation, in the sense that if firm level wages are affected by wages in other firms within the same industry, aggregation from firm to industry level tends to increase the insider weight.

Johansen's results also revealed that the insider weight seems to increase over time, accompanying the increase in the share of the wage drift over the sample period. This evidence seems to support the hypothesis that more decentralized wage setting makes insider forces in wage determination more important. The study also provides evidence in favor of asymmetric insider effects, as wages seem to be more responsive to insider variables in good times and in expanding industries. Finally, no robust evidence was found in favor of insider hysteresis effects in the Norwegian process of wage formation.

Lever and van Werkhooven (1996) examined the impact of internal and external factors in the process of wage formation in 68 Dutch manufacturing industries (3 digit level) over the time period 1974-86. They found that both internal and external forces play an important role in the process of wage determination in Dutch manufacturing industries. The estimate of the weight of internal factors is 12%, while that of external forces is 88%. Internal factors, however, seem to be more important in large than in small firms. In large firms λ is estimated at 17%-20% and in small firms at 0%-1%. Moreover, product market power (measured by market concentration) appears to exert a significantly positive impact on the wage rate. After including an interaction term between the concentration ratio and nominal productivity, the estimates seem to reveal that the weight of internal factors is higher in more concentrated industries, whereas the direct impact of market concentration on wages disappears. The change in employment appears to have a negative impact on wages, and so no evidence

was reported in favor of insider or hysteresis effects.

Finally, Teulings and Hartog (1998), following the approach of Holmlund and Zetterberg (1991), estimated the long-run wage elasticity (λ) for Japan, Canada, Australia and the Netherlands. The results obtained for the long-run elasticities for industry price and productivity effects (including the ones obtained by Holmlund and Zetterberg) revealed the strongest effects for decentralized countries and the weakest effects for corporatist countries. The only exception is Japan, whose position is rather sensitive to the inclusion of time dummies in the wage regressions. Indeed, when the aggregate variables are replaced by year dummies the long-run elasticity for industry price and productivity effects in Japan rises from 3% to 17% and from 7% to 20%, respectively. This change leads Japan to a position similar to the non-corporatist countries. In Canada and Australia the industry-specific effects (labor productivity trend and industry-relative price) are large and significant. In the Netherlands, trend productivity has a moderate impact on wages and the price effect is almost absent.

In Table 2.1 the insider weight estimates ($\widehat{\lambda}$) obtained for the various countries in the studies mentioned above are summarized.

Table 2.1: Estimates of λ, the long-run elasticity of firm (industry) wages with respect to firm (industry) revenue per employee

	λ	Country
Firm-level data (manufacturing)		
Nickell and Wadhwani (1990)	0.08-0.15	U.K. (1975-82; 1972-86) (219 firms)
Nickell *et al.* (1994)	0.15-016	U.K. (1975-86) (814 firms)
Bentolila and Dolado (1994)	0.11	Spain (1985-88) (1167 firms)
Forslund (1994)	0.05-0.07	Sweden (1984-88) (128 firms)
Wulfsberg (1997)	0.07	Norway (1976-88) (7323 firms)
Industry-level data		
Holmlund and Zetterberg (1991)	0.07-0.12	Sweden (1965-85)
	0.00-0.03	(28 industries)
	0.03-0.04	Norway (1965-82)
	0.03-0.04	(27 industries)
	0.00-0.01	Finland (1965-85)
	0.00-0.00	(28 industries)
	0.12-0.15	Germany (1965-85)
	0.04-0.10	(25 industries)
	0.48-0.49	U.S. (1965-85)
	0.30-0.38	(28 industries)
Teulings and Hartog (1998)	0.10-0.11	Netherlands (1965-85)
	0.04-0.04	(13 industries)
	0.07-0.20	Japan (1970-80)
	0.03-0.17	(25 industries)
	0.33-0.38	Canada (1972-85)
	0.22-0.25	(27 industries)
	0.20-0.22	Australia (1975-85)
	0.19-0.23	(21 industries)
Nickell and Kong (1992)	0.02-0.50*	U.K. (1961-85)
	*λ for each industry	(14 industries)
Johansen (1996)	0.16-0.25	Norway (1966-87) (117 industries)
Lever and van Werkhooven (1996)	0.12-0.15 (all firms)	Netherlands (1974-86)
	0.197 (large firms)	(68 industries)
	0.007 (small firms)	

Note: In Holmlund and Zetterberg (1991) and Teulings and Hartog (1998), for each country, the first range for λ refers to trend productivity and the second to industry relative price.

4 Data and Estimation Method

4.1 The Data

Our basic data source is the *Social Audit* ("Balanço Social") and includes a panel of Portuguese firms having at least 100 employees from all sectors, over the period 1993-97.

The *Social Audit* is gathered annually by the Portuguese Ministry of Employment. When it was first introduced (1986) it covered state-owned firms only. Since then its coverage has been spreading, first to firms with at least 500 employees, and since 1992 to firms with at least 100 employees. Responding to this survey is mandatory. On average, 2,050 firms responded to the survey each year, corresponding to a total of 762,000 workers. In fact, the *Social Audit* is characterized by a very high degree of coverage of large firms in Portugal.

Each year, a respondent firm reports data on a large variety of topics concerning the workforce composition and labor costs. This is organized in six major areas: (i) firm's characteristics; (ii) employment; (iii) labor costs; (iv) occupational safety; (v) vocational training; and (vi) fringe benefits.

The collection of firm's characteristics includes information about location, economic activity (SIC codes), legal setting, employment, number of establishments and production (value added). The existence of a unique identification number for each firm allows us to create a longitudinal panel of firms.

The employment block, which is the largest in this survey, collects detailed information about workers' attributes. This includes information about gender, age, skills, schooling, tenure, hours of work, etc. Total employment is also decomposed by type of contract and skill level, which allows one to compute the number of permanent and temporary workers in the end-of-year count.

The information about labor costs includes annual base-wage, regular paid benefits and bonus, irregular benefits and bonus, costs with vocational training, and other fringe benefits. Information about wage dispersion is also available.

One of the main advantages of this data set, besides its coverage and longitudinal nature, is the availability of information on both firm's and workers' characteristiscs. The possibility of controlling the skill composition of the workforce over the years as well as the possibility of computing workers' flows constitutes an important advantage of this data set. The information about employment by type of contract is equally important.

In order to complement the information available in the *Social Audit,* we will also use the data contained on *Quadros de Pessoal* (*QP*). *QP* is an annual mandatory employment survey collected by the Portuguese Ministry of Employment that covers virtually all establishments with wage earners.[10] In each year every establishment with wage earners is legally obliged to fill in a standardized questionnaire. Reported data cover the establishment itself (location, economic activity and employment), the firm (location, economic activity, employment, sales and legal framework) and each of its workers (gender, age, education, skill,

[10] Public administration and domestic servants are excluded.

occupation, tenure, earnings and duration of work). Currently, the data set collects information on around 200,000 firms and 2 million employees.

The information from *QP* about wages will be used in order to compute the outside wage. There are two main reasons to believe that *QP* can provide a reliable measure of the outside wage. The first is its coverage and reliability. By law, the questionnaire is made available to every worker in a public space of the establishment. This requirement facilitates the work of the services of the Ministry of Employment that monitor compliance of firms with the law (e. g., illegal work). Indeed, the administrative nature of the data and its public availability imply a high degree of coverage and reliability. The second is that the information on earnings is very complete. It includes the base wages (gross pay for normal hours of work), seniority payments, regular benefits, irregular benefits and overtime pay.

The outside wage per employee will be defined by region (NUTs III),[11] excluding the firm's own wage.

Since the data on value-added are not available in the *Social Audit,* the information on sales from *QP* will be used instead. This is possible because the identification code of firms in the *Social Audit* and the *QP* data sets is the same.

The sales variable will be used to compute a measure of productivity and a measure of market share. In fact, the annual sales per employee will be used as a proxy for productivity.[12] The market share is obtained by the ratio between firm's sales and total (5 digit) sector's sales.[13]

Neither the *Social Audit* nor the *QP* data set have information on union density. In fact, there are no micro-data in Portugal regarding the number of workers who are members of a trade union. In order to overcome this difficulty, some proxies will be used to measure workers' bargaining power.

In order to control for the aggregate outside labor market conditions, we decided to include a set of time dummies and the regional unemployment rate. The regional unemployment rate is defined at the level of NUTs II.[14]

A precise definition of all variables is presented in Appendix A.

The sample was limited to firms for which data are held for at least four consecutive years and with no missing values in the explanatory variables. To minimize the effects of outliers, we also excluded from the sample those firms whose real sales increase more than 5 times or decrease to less than one-fifth from one year to the next.[15]

After these restrictions, we obtained an unbalanced panel of 896 firms and 4,391 observations. In table 2.2 some selected variables are reported in order to characterize the sample over the 1993-97 period.

[11] At NUTs III mainland Portugal is split into 28 geographical areas.

[12] It should be noted that in each year information on sales lagged by one and two years is also available, which enables us to instrument sales with two more lags than the other variables.

[13] According to the Portuguese Classification of Economic Activities (CAE), there are 448 sectors at the 5 digit level.

[14] At NUTs II mainland Portugal is split into 5 geographical areas.

[15] This condition reduced the sample by 70 firms.

Table 2.2: Basic Characteristics of the Data
All Firms (1993-97)

	1993	1994	1995	1996	1997
Variables					
Employment	320695	330125	329360	332146	329843
Permanent employment	265457	270871	270990	269643	265954
Employment growth rate	-0.027	-0.005	-0.004	-0.010	0.004
Gross worker turnover	0.085	0.080	0.081	0.079	0.082
Accession rate	0.040	0.040	0.040	0.039	0.041
Separation rate	0.046	0.041	0.041	0.041	0.040
Market Share	0.079	0.096	0.093	0.091	0.094
Firm size	380.0	368.4	367.6	370.7	384.0
Sales per employee*	11620.6	12130.2	12405.9	12642.0	13882.1
Wages per employee*	1178.6	1188.6	1221.9	1251.7	1306.4
Outside wage per employee*	874.7	913.2	902.8	929.2	954.2
Number of firms	844	896	896	896	859

Note: * Annual real values [in PTE (thousand escudos); 1 EUR≡200.482 PTE]; CPI deflator (base=1991).

As can be seen in Table 2.2, the firms in the sample have an average size of 374 employees, 9.1% of market share, representing a total of about 328500 workers. Permanent employment represents around 82% of total employment. Over the 1993-97 period, employment growth rates changed between a negative value of 2.7% in 1993 to a positive one of 0.4% in 1997, representing total workers' accessions and separations of 8.1% of average employment.

Between 1993 and 1997 real average wages in the firm grew, on average, at an annual rate of 2.1%, whereas the outside wage in the region grew at an annual rate of 1.8%. The average annual growth rate of real sales per employee was 3.6%.

4.2 Estimation Method

The dynamic linear model of equation (2.12) is an autoregressive fixed effects model. In the presence of a model of such nature the ordinary least squares (OLS) estimator is inconsistent, due to the existence of correlation between the lagged endogenous variable and the disturbances. A conventional way to tackle this problem is to use an instrumental variables estimation method. The application of the generalized method of moments (GMM) estimator suggested by Arellano and Bond (1991) overcomes these difficulties, producing consistent estimates. The GMM estimator identifies the parameters of the model under the assumption of lack of serial correlation in the error terms, and as this assumption is essential for the consistency of the estimator, a test of autocorrelation, developed in Arellano and Bond (1991), will also be reported.

The empirical model will be estimated using the system (SYS) GMM estimator proposed by Arellano and Bover (1995) and Blundell and Bond (1998). The SYS GMM estimator uses lagged first differences as instruments for equations in levels in addition to the usual lagged levels as instruments for equations in first-differences. This option is justified by the fact that the SYS GMM estimator can dramatically improve the performance of the traditional first-differences (DIF) GMM estimator when the autoregressive parameter is moderately high and the number of time-series observations is moderately small. Indeed, the SYS GMM estimator has superior properties in terms of small sample bias and root mean squared error, especially for persistent series [see Blundell and Bond (1998)].[16]

In fact, recent empirical studies have reported some problems with the estimation of dynamic panel data models using the DIF GMM estimator in cases of highly persistent regressors, which imply weak correlation of lagged levels with subsequent first differences. Mairesse and Hall (1996) and Blundell and Bond (1998), for example, showed that when the panel data are characterized by a large sample of firms observed over a small number of time periods, standard GMM estimators, which eliminate unobserved firm-specific effects first-differencing, have been found to produce unsatisfactory results. Blundell and Bond (2000) and Blundell *et al.* (2000) applied the SYS GMM estimator to panel production functions for the U.S. and also showed that the use of the SYS GMM estimator not only greatly improves the precision of the regression coefficient estimates but also greatly reduces the finite sample bias.

5 Empirical Results

5.1 Measuring the Weight of Insider Forces

The SYS GMM estimates of equation (2.12) for the unbalanced panel of 896 firms from all industries for the period 1994-97 are displayed in column 1 of Table 2.3.[17] In order to control for the skill composition of the firm's workforce, the specification of column 1 includes a set of controls for workers' skills. Thus, five levels of education (omitted category is basic school and less than basic school) and six levels of qualifications (omitted category is apprentices) were added to equation 2.12.[18]

[16] The empirical model was first estimated using the DIF GMM estimator, but unreliable estimates were obtained for the wage equation coefficients.

[17] The equations are estimated using DPD98 (Dynamic Panel Data software) written by Arellano and Bond (1998).

[18] The Wald test of joint significance of the education and qualification levels rejects the hypothesis that the coefficients are all equal to zero. Moreover, the results revealed that controlling for workers' skills reduces the effect of nominal productivity on wages by around 5 percentage points, suggesting that there might exist a positive correlation between workers' skills and nominal productivity.

The results of column 1 report a value of the insider weight (λ) of 22%, estimated with precision.[19] This value is considerably higher than those obtained for other European Countries such as Spain and the U.K. using firm-level data (see Table 2.1). In fact, the effect of nominal productivity on wages is strong and significant (coefficient estimate of 0.17), suggesting that in Portugal wages are highly responsive to firm's performance.

Other evidence is uncovered by the results. First, market share exerts a positive and significant impact on wages, suggesting that monopoly power generates monopoly rents that are captured by the employees in the form of higher wages.

Second, as expected, the regional unemployment rate has a negative and significant impact on wages. The elasticity of the regional unemployment rate is -0.115, which is a value that is in accordance with previous estimates [see, for instance, Blanchflower *et al.* (1990) and Nickell *et al.* 1994]. This result reveals that outsiders' forces have an important role in wage determination in the sense that they affect the alternative options to the bargaining parties.

Finally, there is no evidence of membership hysteresis effect. In fact, the coefficient on the employment change (Δn) is negative and not statistically different from zero. It should be noted that employment growth raises wages because, given current membership, the lower the last period's employment (i. e. the higher is employment growth) the more protected from losing their jobs the insiders will be.

The insignificant impact of employment changes on bargained wages is not unexpected and, indeed, previous empirical research has found little support for the hypothesis that current wages depend inversely on past employment. The only exception is the result obtained by Bentolila and Dolado (1994), who found robust evidence in favor of insider hysteresis effects, when the insiders are measured by the number of permanent employees. We test this hypothesis, measuring the membership effect by the change in permanent employment rather than in total employment, but the coefficient on the employment change remained imprecisely estimated.

In order to compare with previous studies using manufacturing data, the same regression was reestimated for the manufacturing sector. In the sample, manufacturing firms represent about 61% of the total of firms. The results for the manufacturing sector are presented in column 2 of Table 2.3.

The results obtained for the sample of manufacturing firms are, in general, quite similar to those obtained for the full sample, although the insider effect appears to be rather weak (λ=0.18). The only exception concerns the impact of employment growth on wages. In fact, in manufacturing firms the employment growth has a significant and positive impact on wages (see column 2 of Table 2.3), in contrast to the negative effect obtained in the overall sample.

It should be noted that in the two regressions of Table 2.3, the test statistics reported in both cases verify the critical assumption of no second-order serial

[19] The long-run value of the insider weight is calculated by dividing the nominal productivity (the short-run coefficient) by one minus the coefficient on the lagged wages.

correlation (m_2 test) and the validity of the instruments (Sargan test) at the conventional levels of significance.

Table 2.3: SYS GMM Estimates of Wage Equation (1994-97)
Measuring the Weight of Insider Forces
Dependent variable: Wages $(w_{it} - \overline{w}_{jt})$

	(1) All Firms	(2) Manufacturing Firms
Independent Variables		
Wages lagged $(w_{it-1} - \overline{w}_{jt})$	0.230**	0.160
	(2.0)	(1.1)
Nominal productivity $(p_{it} + y_{it} - n_{it} - \overline{w}_{jt})$	0.170*	0.148*
	(4.2)	(3.3)
Employment growth (Δn_{it})	-0.285	0.323**
	(-1.2)	(2.1)
Market share (MS_{it-1})	0.017*	0.018*
	(3.0)	(3.1)
Regional unemployment rate (u_{jt-1})	-0.115*	-0.106*
	(-4.0)	(-3.5)
Education Levels		
Preparatory and lower secondary	-0.007	0.002
	(-1.1)	(0.2)
Upper secondary	0.007	0.015**
	(1.0)	(2.3)
College	0.057*	0.064*
	(5.0)	(5.3)
Others	-0.001	-0.014***
	(-0.3)	(-1.7)
Qualification Levels		
Manager and highly Professional	0.012	0.010
	(0.9)	(0.7)
Professional	0.015**	0.031*
	(2.2)	(3.4)
Supervisors	-0.001	0.011
	(-0.1)	(1.4)
Highly skilled and unskilled	0.011	0.001
	(1.4)	(0.1)
Semi-skilled and unskilled	-0.006	-0.004
	(-1.1)	(-0.7)
Constant	-0.200	-0.023
	(-1.3)	(-0.2)
Time dummies	yes*	yes*

Table 2.3: Continued

Wald (p-value)	2047.8 (.000)	1148.1 (.000)
Sargan (p-value)	33.1 (.232)	33.6 (.215)
m_1(p-value)	-3.0 (.002)	-2.0 (.045)
m_2(p-value)	1.6 (.101)	0.7 (.482)
NT	3495	2138
λ (insider weigt)	0.22	0.18

Notes: (i) Subscript i denotes firms, j refers to region and t denotes time;
(ii) t-statistics in parentheses;
(iii) *, **, *** denote significant, at 1, 5 and 10 percent, respectively;
(iv) The variables treated as endogenous are: $(w-\overline{w})$, Δn and $(p+y-n-\overline{w})$; instruments used are: $(w-\overline{w})_{(-2)}...(w-\overline{w})_{(-4)}$, $\Delta n_{(-2)}...\Delta n_{(-4)}$, $(p+y-n-\overline{w})_{(-2)}...(p+y-n-\overline{w})_{(-6)}$, $\Delta(w-\overline{w})_{(-1)}...\Delta(w-\overline{w})_{(-3)}$, $\Delta(\Delta n)_{(-1)}...\Delta(\Delta n)_{(-3)}$ and $\Delta(p+y-n-\overline{w})_{(-1)}...\Delta(p+y-n-\overline{w})_{(-5)}$.

5.2 Measuring Insider Power

The results obtained so far were not robust concerning the importance of insider power in wage determination. In fact, regarding the membership effect associated with Δn, the estimates reported in Table 2.3 revealed a negative (and not statistically significant) coefficient for the full sample in contrast with a positive (and statistically significant) effect for the sample of manufacturing firms. Even though the data set has no information about union density, in this Section we use a set of proxies of workers' bargaining power in order to examine the extent to which insider power has an important role in wage determination. As pointed out by Lindbeck and Snower (2002), "...the insider-outsider theory is not just about labor unions. Any employee whose position is protected by labor turnover costs is an insider of sorts, regardless of whether he belongs to a union". Since Portugal is characterized by a stricter employment legislation with higher firing costs and low flows in and out of employment, it appears that there is some scope for the existence of insider power beyond the one that might result from the behavior of unions.

The first measure used as a proxy of workers' bargaining power is the lay-off rate. The lay-off rate ($layoff$) is measured as the ratio between the total number of separations in the year and the firm's average employment in the year (in logs). In order to minimize the endogenity problems the variable is lagged by one year.

When wages are largely set in the interest of the insiders workers, as predicted by the insider-outsider theory, it should be expected that lay-off rates have a negative impact on wages, since higher lay-off rates threaten the jobs of the insiders. Graafland (1992) using aggregate data for the Netherlands, showed that the lay-off rate has a significant negative influence on wages.

The second measure of workers' bargaining power is defined as the proportion of permanent employees in the firm in the end-of-year count (*pper*). According to the framework developed by Bentolila and Dolado (1994), the proportion of temporary employment might affect wages in at least two ways. One is the *buffer effect*, according to which when the proportion of temporary employees rises the union's employment target becomes smaller relative to expected total employment, and insiders' representatives ask for higher wages. The second is the *bargaining effect*, which comprises two sub-effects working in opposite directions. The *harassment effect* (Lindbeck and Snower, 1988) has a positive impact on wages. The idea is that insiders strengthen their bargaining position by threatening to be uncooperative with temporary employees. If by becoming relatively less numerous insiders become more essential to the production process, the firm will prefer to accept their wage demands rather than keep permanent workers' positions vacant or replace them with new workers with lower productivity. The *discipline effect* acts in the opposite direction. As the proportion of temporary workers rises, strike activity declines (given their lower firing costs, temporary workers are more reluctant to engage in strikes), which erodes the power of permanent workers and consequently the capacity to extract rents. Besides these two bargaining effects observed by Bentolila and Dolado (1994), there might be another one that can be called the *specific-training effect*. As long as labor turnover costs (namely, training costs) rise with job tenure, the proportion of permanent employees may be used as a measure of the specific training asset of the firm. Moreover, investment in firm-specific human capital is also expected to be encouraged by high labor turnover costs. Thus, it will be expected that as the proportion of permanent workers increases, insiders are in a better position to extract rents, since it becomes more costly to the firm to replace their high-productivity permanent workers with new low-productivity workers.

The third proxy for workers' bargaining power is a measure of the labor utilization rate within the firm. As initially suggested by Gregory (1986), it is probably the labor utilization rate within the firm that is particularly important for wage negotiations, rather than the labor utilization rate within the economy. The labor utilization rate may affect wages in two ways. First, higher labor utilization rates within the firm increases the probability of job retention of an insider worker raising their power of negotiation. Second, as the labor utilization rate increases, the threat of a strike becomes more credible to the firm, raising the workers' bargaining power.

The labor utilization rate (*lur*) is defined as the ratio between the total number of hours actually worked in the year and the maximum annual potential of hours worked (in logs and lagged by one year). Higher labor utilization rates within the firm will induce workers to demand higher wages, *ceteris paribus*.

Table 2.4 reports the SYS GMM estimates of the basic specification presented in Table 2.3 with the inclusion of various measures of workers' bargaining power. Although not reported, these regressions include the same set of controls

for workers' skills (five educational levels and six qualification levels).[20] [21]

In column 1 of Table 2.4 the lay-off rate is included. As expected, the lay-off rate has a negative and significant impact on wages (-0.019), suggesting that when the employment perspectives of employed workers worsen, they tend to restrain wage demands.

In column 2 of Table 2.4 the proportion of permanent employees in total employment is included. The parameter estimate of the proportion of permanent employees reports a positive and significant impact on wages of 0.093. This evidence seems to suggest that a positive bargaining effect dominates in firms with a higher proportion of permanent workers.[22] Workers employed in firms with a more stable workforce are in a better position to extract rents in the form of higher wages. This result is consistent with one of the predictions of the insider-outsider theory, that wages will be relatively high in sectors (firms) with high labor turnover costs and where insiders enjoy relatively high job security.

In column 3 of Table 2.4 the labor utilization rate within the firm was introduced as a measure of workers' bargaining power. The coefficient of the utilization rate has the expected sign (0.132) and is statistically significant. Hence, workers in firms with higher labor utilization rates have higher insider power and, thus, earn more.

Finally, in column 4 of Table 2.4 we decided to include all the three measures of workers' bargaining power. When the three measures are simultaneously included, the lay-off rate becomes statistically insignificantly different from zero. The impact on wages of the proportion of permanent employees and the labor utilization rate remain positive and significant, suggesting that insider power has an important role in the process of wage determination in Portugal. It should be noted that the coefficients of the other variables remain pratically unchanged with the introduction of a set of controls for workers' bargaining power.

On balance, the results presented in this section show that the insider-outsider considerations have empirical relevance for Portugal, a country characterized by the existence of high labor turnover costs, low worker and job flows and high unemployment duration. Indeed, the results reveal that firms where insider workers have more market power tend to pay higher wages. In particular, in firms with low lay-off rates, a high proportion of permanent employees and high rates of labor utilization within the firm, workers seem to extract rents in the form of higher wages.

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[20] In all regressions of Table 2.4 the test statistics reported in all cases verify the critical assumption of no second-order serial correlation (m_2 test) and the validity of the instruments (Sargan test) at the conventional levels of significance.

[21] These same results are reported in Table B.2.1 of Appendix B for the manufacturing sector.

[22] This result holds even when we control for workers' tenure on the job, suggesting that the positive effect of the proportion of permanent employees on wages is not due to the presence of a composition effect. A coefficient estimate of 0.067 was obtained for the variable proportion of permanent workers when a control for workers' tenure was included in the model.

Table 2.4: SYS GMM Estimates of Wage Equations - All Firms (1994-97)
Measuring Insider Power
Dependent variable: Wages $(w_{it} - \overline{w}_{jt})$

	(1) Lay-off Rate	(2) Proportion of Permanent Employees	(3) Labor Utilization Rate	(4) All Measures
Independent Variables				
Wages lagged $(w_{it-1} - \overline{w}_{jt})$	0.173	0.233**	0.231**	0.199***
	(1.5)	(2.1)	(2.0)	(1.8)
Nominal productivity $(p_{it} + y_{it} - n_{it} - \overline{w}_{jt})$	0.184*	0.178*	0.169*	0.181*
	(4.5)	(4.3)	(4.2)	(4.4)
Employment growth (Δn_{it})	-0.282	-0.334	-0.300	-0.312
	(-1.3)	(-1.5)	(-1.3)	(-1.5)
Market share (MS_{it-1})	0.018*	0.014*	0.017*	0.015*
	(3.1)	(2.7)	(3.0)	(2.8)
Regional unemployment rate (u_{jt-1})	-0.118*	-0.104*	-0.116*	-0.109*
	(4.1)	(-3.9)	(-4.0)	(-4.0)
Lay-off rate $(layoff_{it-1})$	-0.019*			-0.010
	(-2.6)			(-1.4)
Proportion of permanent employees $(pper_{it-1})$		0.093*		0.093*
		(4.7)		(4.8)
Labor utilization rate (lur_{it-1})			0.132**	0.153*
			(2.3)	(2.7)
Constant	-0.266***	-0.217	-0.205	-0.236
	(-1.8)	(-1.4)	(-1.3)	(-1.6)
Time dummies	yes*	yes*	yes*	yes*
Wald (p-value)	2002.0 (.000)	2263.9 (.000)	2129.7 (.000)	2246.0 (.000)
Sargan (p-value)	30.9 (.323)	34.0 (.201)	34.0 (.202)	33.6 (.214)
m_1(p-value)	-2.9 (.004)	-3.0 (.002)	-3.1 (.002)	-3.0 (.003)
m_2(p-value)	1.4 (.158)	1.7 (.096)	1.7 (.096)	1.6 (.120)
NT	3495	3495	3495	3495
λ (insider weight)	0.22	0.23	0.22	0.23

Notes: see notes to Table 2.3; each regression includes five educational levels and six qualification levels.

5.3 Testing for Asymmetric Insider Effects

Insider effects may be more important in expanding firms when compared to declining ones, and similarly, firms' wages may be more responsive to insider variables in good than in bad times. Such asymmetric insider effects imply downward wage rigidity, and will tend to put more pressure on employment when times are bad. Even though some empirical evidence was found that wage adjustments are asymmetric [see, for instance, Nickell and Wadhwani (1990), Blanchflower (1991), Holzer and Montgomery (1993) and Johansen (1996)], this issue remains unsettled and further evidence seems to be required. In this section we test the extent to which wages in Portugal are more responsive to insider variables in the face of rising demand than in the face of declining demand.

The major difficulty associated with the implementation of any test of asymmetry is that demand is not observed. In order to have a measure of expected demand we used the average rate of growth of real sales over the last three years as a proxy. Then we interact a dummy that takes the value one for positive rates of sales growth in the last three years (zero otherwise) with some of the insider variables. Sales growth is treated as endogenous and thus all interaction terms are also treated as endogenously determined. The regression estimates of a specification that includes just an interaction term with nominal productivity are reported in column 1 of Table 2.5.[23]

As can be seen from Table 2.5, the interaction term between the average rate of sales growth dummy and nominal productivity is not statistically different from zero, suggesting that wage adjustments show no consistent pattern of asymmetry to positive and negative shifts in sales. To test the robustness of this hypothesis we decided to interact the insider variables with employment changes. Blanchflower (1991) showed that workers employed in firms where employment is expected to grow receive a pay premium while those employed in contracting plants suffer no pay disadvantage.

Thus, a dummy taking the value one when $\Delta n_{it} > 0$ (zero otherwise) was interacted with nominal productivity. Since employment growth is endogenous, the interaction term is treated as endogenously determined. Once again, the interaction term is not statistically significant, rejecting the hypothesis of the existence of asymmetric insider effects.

In sum, the asymmetry test seems to reveal that in Portugal wages are flexible in both directions: upward and downward. Wages are equally responsive to insider variables when demand is expected to rise or to decline. This result reinforces the idea that wages are quite flexible in Portugal.

[23] A regression was estimated including also an interaction term with lagged wages but the results remained pratically unchanged.

Table 2.5: SYS GMM Estimates of Wage Equations - All Firms (1994-97)
Testing for Asymmetric Insider Effects
Dependent variable: Wages $(w_{it} - \overline{w}_{jt})$

	(1)	(2)
Independent Variables		
Wages lagged $(w_{it-1} - \overline{w}_{jt})$	0.262*	0.186**
	(2.7)	(2.1)
Nominal productivity $(p_{it} + y_{it} - n_{it} - \overline{w}_{jt})$	0.151*	0.161*
	(4.3)	(4.9)
Employment growth (Δn_{it})	-0.044	-0.165
	(-0.3)	(-0.9)
Market share (MS_{it-1})	0.018*	0.020*
	(3.9)	(4.2)
Regional unemployment rate (u_{jt-1})	-0.115*	-0.118*
	(-4.6)	(4.6)
Interaction Terms		
Nominal productivity*Sales growth dummy	-0.003	
	(-0.4)	
Nominal productivity*Employment growth dummy		0.010
		(0.7)
Constant	-0.131	-0.135
	(-1.0)	(-1.0)
Time Dummies	yes*	yes*
Wald (p-value)	2506.9 (.000)	2016.5 (.000)
Sargan (p-value)	46.1 (.385)	44.3 (.461)
m_1(p-value)	-3.0 (.003)	-2.9 (.004)
m_2(p-value)	1.5 (.126)	1.4 (.158)
NT	3495	3495

Notes: see notes to Table 2.3; each regression includes five educational levels and six qualification levels.

6 Conclusion

This study investigates wage determination at the firm level using a longitudinal panel of large firms in Portugal. The main empirical findings are the following.

First, insider forces such as revenue per employee and market share have a significant impact on wage determination. This is true for the overall sectors and, in particular, for manufacturing. After controlling for the skill mix of the workforce, the full sample estimates imply a long-run insider weight of around 22%-23% (between 17% and 19% for the manufacturing sector), which is comparable with estimates reported for economies characterized by a decentralized system of wage negotiation. The system of wage bargaining in Portugal exhibits characteristics of a centralized and regulated industrial regulation system, nevertheless the results of this study reveal that there is room for firm maneuvering. In fact, and even though collective bargaining imposes a binding constraint on the employer wage policy, wages seem to be very sensitive to firm-specific conditions, i. e, firm wages in Portugal are significantly affected by firm performance.

Second, outside labor market conditions measured by the regional unemployment rate also play an important role in wage determination. The regional unemployment level in the economy has an influence on the negotiated wage through the probabilities of finding a job. Thus, the negative and significant impact of the regional unemployment rate on wages suggests that workers are more inclined to accept wage moderation when the probabilities of finding a job worsen. The results also revealed that workers restrain wages when the labor utilization rate within the firm declines.

Third, the results presented in this study also showed that the insider-outsider considerations have significant empirical relevance for Portugal, a country characterized by the existence of high labor turnover costs, low flows into and out of employment and high unemployment duration. More specifically, the results indicated that firms where insider workers seem to have more market power tend to pay higher wages.

When the conventional measure to evaluate insider power was used, namely the change in the number of insiders, no robust evidence was found that insider power has an important role in wage determination. However, when other measures of insider power were included in the wage equation model, empirical evidence was found that is consistent with the idea that wages will be higher in sectors (firms) with high labor turnover costs and high job retention probabilities. In particular, in firms with low lay-off rates, a high proportion of permanent workers and high rates of labor utilization, insider workers are in a better position to extract rents in the form of higher wages.

Finally, no robust evidence was found in favor of the existence of asymmetric insider effects, contradicting some previous empirical evidence that wage adjustments are asymmetric. Wages in Portugal seem to be flexible in both directions: upward and downward.

APPENDIX A - Variables: Definition[24] and Source[25]

Average employment: defined as the mean between the number of workers at the beginning of the year and the number of workers at the end of the year; *Social Audit*;

Wages: annual real labor cost (base wage + regular paid benefits and premiums) divided by average employment; *Social Audit*;

Nominal productivity: annual sales at constant prices divided by average employment; *Quadros de Pessoal* and *Social Audit*;

Market share: total sales in each firm divided by total sales in the sector defined at 5 digits according to CAE; *Quadros de Pessoal*;

Employment growth: measured as the annual rate change in the total number of workers; *Social Audit*;

Gross Worker Turnover: total number of workers who joined (accessions) and left (separations) the firm over the year divided by average employment; *Social Audit*;

Accession rate: total number of workers who joined (accessions) the firm over the year divided by average employment; *Social Audit*;

Separation rate≡Lay-off rate: total number of workers who left (separations) the firm over the year divided by average employment; *Social Audit*;

Labor utilization rate: the ratio between the number of total hours actually worked and the maximum annual potential of worked hours; *Social Audit;*

Proportion of permanent employees: the number of permanent employees in the end-of-year count as a proportion of total employment in the end-of-year count; *Social Audit*;

Outside wage: aggregate real wage by region (defined at NUTs III) excluding the firm's own wage (per employee); *Quadros de Pessoal*;

Regional unemployment rate: defined at the level of NUTs II; *Employment Survey - INE (Instituto Nacional de Estatística)*;

Education: five educational levels were defined (proportion of workers) - primary and less than primary (the omitted category), preparatory and lower secondary, upper secondary, college and others; *Social Audit*;

[24] All variables are in logs.
[25] Data sources are in italics.

Qualification: six qualification levels were defined (proportion of workers) - manager and highly professional, professional, supervisors, skilled and highly skilled, semiskilled and unskilled and apprentices (the omitted category); *Social Audit;*

Price deflator: Consumer Price Index (1991=100); *Consumer Price Index - INE.*

APPENDIX B - SYS GMM Estimates - Manufacturing Firms

Table B.2.1: SYS GMM Estimates of Wage Equations - Manufacturing Firms (1994-97)
Measuring Insider Power
Dependent variable: Wages ($w_{it} - \overline{w}_{jt}$)

	(1) Lay-off Rate	(2) Proportion of Permanent Employees	(3) Labor Utilization Rate	(4) All Measures
Independent Variables				
Wages lagged ($w_{it-1} - \overline{w}_{jt}$)	-0.042	0.177	0.163	0.026
	(-0.3)	(1.2)	(1.1)	(0.2)
Nominal productivity ($p_{it} + y_{it} - n_{it} - \overline{w}_{jt}$)	0.201*	0.146*	0.143*	0.181*
	(4.3)	(3.3)	(3.2)	(4.0)
Employment growth (Δn_{it})	0.320**	0.307***	0.328**	0.317**
	(2.2)	(1.9)	(2.2)	(2.1)
Market share (MS_{it-1})	0.020*	0.018*	0.018*	0.019*
	(2.9)	(3.1)	(3.1)	(2.9)
Regional unemployment rate (u_{jt-1})	-0.128*	-0.104*	-0.106*	-0.120*
	(-3.9)	(-3.4)	(-3.4)	(-3.7)
Lay-off rate ($layoff_{it-1}$)	-0.020*			-0.015**
	(-2.8)			(-2.2)
Proportion of permanent employees ($pper_{it-1}$)		0.082*		0.088*
		(2.9)		(3.0)
Labor utilization rate (lur_{it-1})			0.105	0.119
			(1.5)	(1.6)
Constant	-0.148	-0.017	-0.007	-0.092
	(-0.9)	(-0.1)	(-0.1)	(-0.6)
Time dummies	yes*	yes*	yes*	yes*
Wald (p-value)	871.7 (.000)	1154.6 (.000)	1139.3 (.000)	926.4 (.000)
Sargan (p-value)	36.5 (.130)	33.4 (.223)	34.2 (.195)	36.6 (.128)
m_1(p-value)	-1.6 (.112)	-2.1 (.036)	-2.1 (.039)	-1.9 (.063)
m_2(p-value)	-0.0 (.974)	0.7 (.454)	0.7 (.472)	0.2 (.823)
NT	2138	2138	2138	2138
λ (insider weight)	n.d.	0.18	0.17	0.19

Notes: see notes to Table 2.3; each regression includes five educational levels and six qualification levels.

References

[1] Abowd, J. and T. Lemieux (1993), "The Effects of Product Market Competition on Collective Bargaining Agreements: The Case of Foreign Competition in Canada", *Quarterly Journal of Economics*, Vol. 108, pp. 983-1014.

[2] Arellano, M. and S. Bond (1991), "Some Tests of Specification for Panel Data: Monte Carlo Evidence and an Application to Employment Equations", *Review of Economic Studies*, Vol. 58, pp. 277-297.

[3] Arellano, M. and S. Bond (1998), "Dynamic Panel Data Estimation using DPD98 for Gauss: a Guide for Users", *mimeo.*

[4] Arellano, M. and O. Bover (1995), "Another Look at the Instrumental Variable Estimation of Error-components Models", *Journal of Econometrics*, Vol. 68, pp. 29-51.

[5] Bentolila, S. and J. Dolado (1994), "Labour Flexibility and Wages: Lessons from Spain", *Economic Policy*, Vol. 18, pp. 55-99.

[6] Blanchard, O. (1991), "Wage Bargaining and Unemployment Persistence", *Journal of Money, Credit, and Banking*, Vol. 23, pp. 277-292.

[7] Blanchard, O. and P. Diamond (1994), "Ranking, Unemployment Duration and Wages", *Review of Economic Studies*, Vol.61, pp. 417-434.

[8] Blanchard, O. and P. Portugal (2001), "What Hides Behind an Unemployment Rate: Comparing Portuguese and U.S. Labor Markets", *American Economic Review*, Vol. 91, pp. 187-207.

[9] Blanchard, O. and L. Summers (1986), "Hysteresis and the European Problem", *NBER Macroeconomics Annual*, Cambridge, MIT Press.

[10] Blanchflower, D. (1991), "Fear, Unemployment and Pay Flexibility", *Economic Journal*, Vol. 101, pp. 483-496.

[11] Blanchflower, D., A. Oswald and M. Garrett (1990), "Insider Power in Wage Determination", *Economica*, Vol. 57, pp. 143-170.

[12] Blanchflower, D., A. Oswald and P. Sanfey (1996) "Wages, Profits and Rent-Sharing", *Quarterly Journal of Economics*, Vol. 111, pp. 227-251.

[13] Blundell, R. and S. Bond (1998), "Initial Conditions and Moment Restrictions in Dynamic Panel Data Models", *Journal of Econometrics*, Vol. 87, pp. 115-143.

[14] Blundell, R. and S. Bond (2000), "GMM Estimation with Persistent Panel Data: an Application to Production Functions", *Econometric Review*, Vol. 19, pp. 321-340.

[15] Blundell, R., S. Bond and F. Windmeijer (2000), "Estimation in Dynamic Panel Data Models: Improving on the Performance of the Standard GMM Estimators", Institute for Fiscal Studies Working Paper WP 00/12.

[16] Bruno, M. and J. Sachs (1985), *The Economics of Worldwide Stagflation*, Oxford: Basil Blackwell.

[17] Forslund, A. (1994) "Wage Setting at the Firm Level - Insider versus Outsider Forces", *Oxford Economic Papers*, Vol. 46, pp. 245-261.

[18] Gregory, R. (1986) "Wages Policy and Unemployment in Australia", *Economica*, Vol. 53, pp. S53-S74.

[19] Gottfries, N. and H. Horn (1987), "Wage Formation and Unemployment Persistence", *Economic Journal*, Vol. 97, pp. 877-886.

[20] Graafland, J. (1992), "Insiders and Outsiders in Wage Formation: The Dutch Case", *Empirical Economics*, Vol. 17, pp. 583-602.

[21] Hildreth, A. and A. Oswald (1997), "Rent-Sharing and Wages: Evidence from Company and Establishment Panels", *Journal of Labor Economics*, Vol. 15, pp. 318-337.

[22] Holmlund, B. and J. Zetterberg (1991), "Insider Effects in Wage Determination - Evidence from Five Countries", *European Economic Review*, Vol. 35, pp. 1009-1034.

[23] Holzer, H. and E. Montgomery (1993), "Asymmetries and Rigidities in Wage Adjustments by Firms", *Review of Economics and Statistics*, Vol. 75, pp. 397-408.

[24] Johansen (1996), K. "Insider Forces, Asymmetries, and Outsider Inefectiveness: Empirical Evidence for Norwegian Industries 1966-87", *Oxford Economic Papers*, Vol. 48, pp. 89-104.

[25] Layard, R., S. Nickell and R. Jackman (1991), *Unemployment: Macroeconomic Performance and the Labour Market*, Oxford: Oxford University Press.

[26] Lever, M. and J. van Werkhooven (1996) "Insider Power, Market Power, Firm Size and Wages: Evidence from Dutch Manufacturing Industries", *Labour Economics*, Vol. 3, pp. 93-107.

[27] Lindbeck, A. and D. Snower (1985), "Explanations of Unemployment", *Oxford Review of Economic Policy*, Vol. 2, pp. 34-69.

[28] Lindbeck, A. and D. Snower (1986), "Wage Setting, Unemployment, and Insider-Outsider Relations", *American Economic Review Proceedings*, Vol. 76, pp. 235-239.

[29] Lindbeck, A. and D. Snower (1988), "Cooperation, Harassment, and Involuntary Unemployment: an Insider-Outsider Approach", *American Economic Review*, Vol. 78, pp. 167-188.

[30] Lindbeck, A. and D. Snower (2002), "The Insider-Outsider Theory: a Survey", IZA Discussion Paper N° 534.

[31] Mairesse, J. and B. Hall (1996), "Estimating the Productivity of Research and Development in French and US Manufacturing Firms: an Exploration of Simultaneity Issues with GMM Methods", in K. Wagner and B. van Ark eds., *International Productivity Differences and Their Explanations*, Elsevier Science, pp. 283-315.

[32] Nickell, S. (1999), "Product Markets and Labour Markets", *Labour Economics*, Vol. 6, pp. 1-20.

[33] Nickell, S. and P. Kong (1992), "An investigation into the Power of Insiders in Wage Determination", *European Economic Review*, Vol. 36, pp. 1573-1599.

[34] Nickell, S., J. Vainiomaki and S. Wadhwani (1994), "Wages and Product Market Power", *Economica*, Vol. 61, pp. 457-473.

[35] Nickell, S. and S. Wadhwani (1990), "Insider Forces and Wage Determination", *Economic Journal*, Vol. 100, pp. 496-509.

[36] Phelps, E. (1972), *Inflation Policy and Unemployment Theory*, London: Macmillan.

[37] Pissarides, C. (1992), "Loss of Skill During Unemployment and the Persistence of Employment Shocks", *Quarterly Journal of Economics*, Vol. 107, pp. 1371-1391.

[38] Organization for Economic Cooperation and Development (1997), *OECD Employment Outlook*, Paris: OECD.

[39] Organization for Economic Cooperation and Development (1999), *OECD Employment Outlook*, Paris: OECD.

[40] Solow (1985), R. "Insiders and Outsiders in Wage Determination", *Scandinavian Journal of Economics*, Vol. 87, pp. 411-428.

[41] Teulings, C. and J. Hartog (1998), *Corporatism or Competition? Labour Contracts, Institutions and Wage Structures in International Comparison*, Cambridge: Cambridge University Press.

[42] Wulfsberg, F. (1997), "An Application of Wage Bargaining Models to Norwegian Panel Data", *Oxford Economic Papers*, Vol. 49, pp. 419-440.

Chapter 3

Earnings Losses of Displaced Workers: Evidence From a Matched Employer-Employee Data set

1 Introduction

Worker displacement has been the subject of an extensive and growing literature.[1] The costs of job loss in terms of unemployment (incidence and duration) and earnings-changes (magnitude and persistence) have been the most studied aspects of job displacement. It is widely agreed that both unemployment and subsequent wages of displaced workers depend on their search behavior. In this respect it has been argued that advance notification of displacement could avoid some losses, mainly through a reduction in the incidence of joblessness.[2]

During the 1980s a number of empirical studies emerged analyzing workers' post-displacement wages and unemployment in the U.S. [(see, for instance, Podgursky and Swaim (1987), Kruse (1988), Addison and Portugal (1989), Kletzer (1989, 1991)].[3] Basically, these studies provide a snapshot view of short-term earnings losses, defined as the difference between pre- and post-displacement earnings of displaced workers. This literature established a basic stylized fact - high-tenure workers suffered the greatest losses in terms of subsequent wages and time employed.

However, this type of analysis, focusing solely on workers who have been displaced, is likely to underestimate the magnitude of wage losses, since it does not account for the earnings growth that would have occurred in the absence of job loss. A simple comparison of pre- with post-separation earnings for displaced workers is insufficient. Following the studies of Ruhm (1991) and Jacobson *et al.* (1993) the empirical investigation on worker displacement and wages has been examining how the wages of displaced workers vary compared to workers who were not displaced. The emphasis in worker displacement research has shifted from short-term wage losses to long-term dynamics. In fact, in recent years the existence of suitable longitudinal data sets in the U.S. and Europe that match workers and firms enables the comparison of wage patterns for displaced and identical non-displaced workers.

More specifically, two main issues are addressed by the recent empirical literature on worker displacement and wages. The first issue examines the existence

[1] Displacement is usually defined as the permanent and involuntary separation of workers from their jobs without cause (i. e. for economic reasons). Displacements occur when a firm closes or downsizes.

[2] The studies of Addison and Portugal (1987), Swaim and Podgursky (1990), Ruhm (1992, 1994) and Jones and Kuhn (1995) examine the effects of advance notice of displacement on unemployment and/or postdisplacement earnings.

[3] See Hamermesh (1989), Fallick (1996) and Kletzer (1998) for surveys.

of earnings losses (relative to non-displaced workers) prior to and after a displacement, as well as their magnitude. The second, which is directly related with the first, is to analyze the persistence of these losses over the time.

For a variety of data sets and methodologies the studies for the U.S. have established that displaced American workers usually experience short spells of unemployment, but also substantial and persistent reductions in earnings, on the order of 10-25%.[4] These losses begin to occur at least three years before displacement [see Jacobson *et al.* (1993)]. For Europe the evidence is less clear-cut. Some studies have concluded for the existence of significant and large earnings losses, while others have concluded for the existence of reduced wage losses. In one point, however, these studies seem to be in agreement. A displaced worker who takes more than a year to find a new job suffers a large penalty on earnings.

It is to be expected that reemployment wages of displaced workers be lower than the wages of those workers that remain employed. As mentioned by Fallick (1996), there are at least four reasons that can explain this pattern. First, the lost of human capital specific to the firm or industry. To the extent that these skills are non-transferable, their contribution to worker's productivity is permanently lost when a job loss occurs. Second, payments by seniority in order to provide incentives not to shirk may delay higher earnings to the latter part of the career. In this case, a permanent separation reduces lifetime earnings. Third, there is the loss of a high quality job match between the worker and the firm. In fact, some authors claim that standard estimates of the return to job-specific training are biased upward by job match and individual unobserved heterogeneity.[5] A long job tenure may signal a high quality match between the firm and the worker and/or a high ability worker, because more able workers and workers in good jobs are less likely to separate. In this line of research, Addison and Portugal (1989) and Kletzer (1989, 1991) showed that tenure in the pre-displacement job is positively associated with post-displacement earnings, reflecting heterogeneity in worker ability and the transferability of skills. Fourth, to the extent that firm's and/or industry characteristics also play a role in the process of wage determination, a displaced worker may lose some wage premium that he/she was previously receiving such as insider rents, union premiums or efficiency wage differentials.

Finally, a fifth reason should be added if a displaced worker experiences some period of unemployment. It is widely argued that a spell of unemployment causes the depreciation of general or transferable worker skills and that prolonged unemployment accelerates the process of depreciation of human capital [Phelps (1972) and Pissarides (1992)]. On the other side, some authors also argue that a period of unemployment may signal an individual as a low productivity worker, leading employers to pay, at least initially, less than his/her marginal productivity [Lockwood (1991)].

The main goal in this study is to analyze the costs of worker displacement

[4] See Kletzer for a survey (1998).

[5] See, among others, the studies of Abraham and Farber (1987, 1988), Altonji and Shakotko (1987), Topel (1991) and Dustmann and Meghir (1999).

in terms of earnings losses. Do displaced workers in Portugal suffer pre- and post-displacement wage losses? If so, what are the magnitude and persistence of these losses?

Two main objectives will drive the investigation. The first objective is to analyse the long-term impact of a displacement on the earnings evolution of displaced workers in Portugal. For this purpose a rich data set that links employers and employees will be exploited. The recent empirical research on this issue has been based on individual or household data sets with little information on the employers in which the individuals work. The possibility to link workers with their employers constitutes a tremendous advantage of this data set, reinforced by the fact that recent empirical work on wage determination has been showing that employers' characteristics are an important determinant of workers' wages [see, for example, the models of insider versus outsiders forces in wage determination as mentioned in chapter 2]. At this point, a quantile regression analysis will be used in order to determine if earnings evolution varies significantly according to the worker's position in the wage distribution. The second objective is to decompose the earnings losses according to their sources/causes.

The issue of job displacement and wages is relevant for several reasons. For this study the most important is that it may help to shed some light on the mechanism of wage determination. If wages reflect primarily general skills we would not expect considerable wage losses upon reemployment; if, on the contrary, wages reflect a large investment in firm (industry)-specific human capital or a high quality job match, then a substantial drop in earnings will be expected. Moreover, if nonemployment implies some depreciation of the stock of human capital or acts as a signal for employers of low productivity workers, then larger earnings losses would be expected for workers experiencing longer spells of nonemployment.

The remainder of this study proceeds as follows. In Section 2 a brief literature review on worker displacement and wages is presented. In Section 3 the longitudinal data are described. In Section 4 the empirical model and regressions estimates are reported. Section 5 presents the contribution of worker and firm characteristics for the differences in earnings. Some concluding remarks follow in Section 6.

2 Literature Review

The methodology of Jacobson *et al.* (1993) introduced a different approach on the study of worker displacement and earnings losses. Basically, their methodology consists in comparing the earnings changes of displaced workers over a long-term period with the earnings changes that would have occurred if the displaced had not lost their jobs. Since this latter variable is not observable, a comparison group of non-displaced workers is used. Using this methodology, the empirical research on this issue has been growing with the appearance of a considerable number of studies for the U.S. and some European countries such

as Germany, France, Britain, Italy and Estonia. In this Section we present a brief summary of the main results obtained in these empirical studies.

Using administrative data on Pennsylvanian workers and their firms over the 1974-86 period, Jacobson *et al.* (1993) found considerable losses both prior to and after displacement for displaced workers when compared with continuously employed workers. High-tenure displaced workers suffer long-term (quarterly) earnings losses averaging 25% per year. These losses start to appear approximately three years before separation and are substantial even for workers reemployed in similar firms. These findings are widespread across gender, age, local labor market conditions, industry and firm size.

Ruhm (1991) obtained comparable results using the PSID (Panel Study of Income Dynamics) data set. He found that weekly wages of U.S. displaced workers in the 1970s were lower by 16% in the year following displacement and, after four years, still 14% lower than those of their non-displaced counterparts. Displacement is also associated with substantially elevated initial unemployment. However, and contrary to earnings losses that are large and persistent, unemployment effects seem to dissipate after a few years.

Stevens (1997) using a national sample of displaced workers from the PSID (1968-88), also found large and persistent earnings losses. In her analysis special attention is given to the effects of multiple job losses on future earnings. Stevens shows that workers who avoid additional displacements have a stronger recovery on annual earnings and hourly wages. Hourly wages are reduced by 9% six or more years after displacement, but only 4% when controlling for the most recent displacement. Moreover, hourly wages began to decline one year prior to job displacement by around 8%.

Bender *et al.* (1999) investigated worker displacement in France and Germany in the 1980s using administrative records. Contrary to the evidence for the U.S. they found, for both countries, that displacement does not induce large (daily) earnings losses. Instead, large penalties were found for displaced workers who were out of employment for more than a year. In fact, a displaced worker in France that takes more than one year to find a new job earns 5% less relative to other displaced workers who are reemployed within the calendar year following displacement. In Germany this loss is around 19%.

Burda and Mertens (2001) studied the effects of displacement on reemployment earnings of German workers using the German Socioeconomic Panel and the Social Security file (1985-94). They found that, on average, a displaced worker subsequently reemployed does not suffer a significant wage reduction. Indeed, in the year following the displacement a full-time displaced worker experiences a wage growth reduction of only 3.6% compared with a similar group of non-displaced workers. They also concluded that large wage losses are associated with changes of industry, but not of firm. Kouch (2001) also used the German Socioeconomic Panel (1986-96) in order to examine the effects of displacement due to plant closing on annual earnings and unemployment duration. He reported an estimated loss of around 13.5% of pre-displacement earnings in the year of displacement and a loss of only 6.5% two years later. No measurable difference in earnings is reported in the years that precede the displacement.

Arulampalam (2001) used the British Household Panel (1991-97) to examine the effect of an interruption of employment on re-entry wages in Britain. After a spell of unemployment a reemployed individual will face a wage penalty of about 6% on re-entry. After three years this penalty rises to 13.5%. Another important conclusion pointed out by the author is that the first spell of nonemployment causes the largest scar on wages. Gregory and Jukes (2001) also examined the effects of unemployment (incidence and duration) on subsequent earnings of a large sample of British men over the 1984-94 period. They found consistent evidence in favor of an unemployment scarring for British workers, due mainly to the effect of unemployment duration. In fact, unemployment incidence has only a temporary effect on wages that vanishes after two years of continuous reemployment. The effect of unemployment duration, by contrast, is permanent and proportional to the length of the spell. In addition to the average penalty of around 10% resulting from the occurrence of an unemployment spell, a worker that experiences a one-year spell of unemployment will suffer an additional penalty of 10 percentage points on first-year earnings after reemployment.

Rosolia (2002) used the Italian Social Security records (1974-97) to analyze the weekly wage patterns prior to and after job separation of displaced Italian workers. His results reveal that displaced workers experience a wage loss both prior to and following separation, with the magnitude of these losses depending on the duration of the nonemployment spells. Workers experiencing a short nonemployment spell (1-2 quarters) and those staying out of employment for more than a year are the ones who bear the largest losses. Four years after separation their earnings are still around 10% below their expected levels. If workers are immediately reemployed in the quarter after displacement the wage gap is negligible. Some evidence was found that high tenure workers at separation are the ones who lose more in the post-displacement period, suggesting that earnings losses are mainly determined by the destruction of a specific match between the worker and his employer. On the contrary, general human capital does not seem to affect the estimated pattern of earnings losses.

Lehmann *et al.* (2002) using data from the Estonian labor force surveys, examined the costs of job loss in Estonia during the period of transition from a planned to a market economy (1989-99). The authors conclude that costs of displacement in Estonia are mainly associated with the risk of nonemployment which affects a minority of workers who experience long-term nonemployment, rather than with a potential penalty in the re-entry wage. In fact, little evidence was found in favor of any wage change differential between displaced and non-displaced workers.

More recently, Bowlus and Vilhuber (2002) and Lefranc (2002) employed a job search framework to analyze the effects of worker displacement on wage outcomes. Specifically, Bowlus and Vilhuber (2002) analyzed the displacement effects on reemployment wages using a longitudinal data set for several U.S. states covering the 1990s that allows one to identify, at the same firm, displaced workers as well as early-leavers. Adopting a search model of displacement they showed that, other things equal, workers who left a "distressed" firm up to two quarters prior to the displacement have higher reemployment wages than do

those who stayed with the same firm until the displacement.

Finally, Lefranc (2002) analyzed the magnitude and sources of wage losses of displaced workers in France (1990-97) and the U.S. (1981-92) using micro data from labor force surveys. He shows that while the magnitude of the wage losses are very similar in the two countries (around 10-15%), the sources of wage adjustment differ considerably. In France, most of the earnings losses result from the loss of accumulated firm-specific human capital, while in the U.S., earnings losses mostly stem from the loss of search rents on the displacement job.

3 The Data

3.1 Description

The data set of this study comes from *Quadros de Pessoal (QP)* and includes all workers that lost their jobs in 1994, 1995 or 1996 due to firm closure and were present in the *QP* registers in at least one of the three years that precedes the displacement event. Indeed, for reasons of completeness we also included in the sample those workers that left the firm (voluntarily or involuntarily) before closing.

QP is an annual mandatory employment survey collected by the Portuguese Ministry of Employment, and covers virtually all firms employing paid labor in Portugal.[6] The data set includes both firm-specific information [location, industry (SIC codes), legal setting, foreign capital, employment, sales, ownership type] and workforce characteristics (labor earnings, worker qualifications, gender, age, tenure, hours of work, etc.).

Moreover, the survey has a longitudinal dimension which makes it particularly well suited for analyzing the issues of firms' entry and exit. Each firm entering the database is assigned a unique identifying number and the Ministry implements several checks to ensure that a firm that has already reported to the database is not assigned a different identification number. By merging the yearly data sets by this identifier it is possible to pinpoint all firms that have entered and exited economic activity. In particular, an exit from the database should signal a firm that has ceased its activity. This criteria, however, is not entirely accurate, due to the fact that some of the firms may temporarily exit the database. A temporary exit may occur for a number of reasons other than cessation of activity, a very likely reason being that the survey form was not received in the Ministry of Employment before the date when the recording operations were closed. Almost all of these temporary exits last less than two years, but can still cause an identification problem if they occur in the terminal years. In order to account for this problem, the information on the last two years after displacement was used solely to control for temporary exits in the intermediate years.[7] Beyond this control, and to ensure that we are in the

[6] Hence, this source does not cover operated family businesses without wage-earning employee and self-employment. Public administration is also excluded.

[7] The last year for which the survey is available is 1998.

presence of firms' true closures and not mergers or acquisitions, we excluded from the sample those workers that appeared in the database in the period after displacement with a year of admission in the new job less than the year of displacement minus one.[8] These exclusions reduced the sample by around 20%.

It is worth noting that workers also have an identification number based on a transformation of his/her social security number which allows us to follow them over the years and to match workers and their firms.

In sum, there are three reasons that make this survey an appropriate source for analyzing the effects of job displacement on earnings. The first is its representativeness. Each year every establishment with wage earners is legally obliged to file a standardized questionnaire, which, by law, is posted in a public space of the firm. Indeed, the administrative nature of the data and their public availability imply a high degree of coverage and reliability. The second results from its longitudinal dimension that enables one to identify firm closings and to follow workers over the years.[9] The third advantage is to have information at both individual and firm level and to match workers with their employers. In fact, almost all empirical research on job displacement and wages has been based on individual or household data sets with little information on the employers for which the individuals work.

Of course, this data set also has disadvantages. The most important is that it is impossible to know if a worker is really in a situation of nonemployment when he/she does not appear in the *QP* registers. In fact, workers who exit the *QP* database, besides being unemployed or out of the labor force, may be self-employed or employed in the public administration sector. To construct a proxy for the existence of a nonemployment spell, we have used the year of admission in the last job as a criterion to classify the workers. Thus, a worker is classified as having experienced a positive spell of nonemployment if he (she) is employed in year t and is not immediately reemployed in year $t+1$.

As mentioned above, to be included in the sample a worker must be present in the *QP* registers at least in one of the three years that precedes the displacement. Adittionally, and in order to ensure that all workers are employed with the same employer three years before separation, the year of admission in the pre-displacement job should be equal to or less than the year of displacement minus three. Even though this data set includes information on usual hours of work, we decided to include only workers employed full-time in the period before separation. In the period after displacement the individual may be reemployed in a part-time job. We also limited the sample to workers aged between 18 and 59 in the year prior to the firm closing. We have also excluded those observations for which some explanatory variable was not available for a particular year and the

[8] If, for example, a worker's displacement year is 1994 and he (she) appears in the database in the post-displacement period with a year of admission in the new job of 1992 or less, he (she) is excluded from the sample.

It should also be noted that when a worker is present in the *QP* files for more than one time in a given year, the register in the firm in which he had worked a higher number of hours was selected.

[9] Eight spells of *QP*, from 1991 to 1998, were used for this study.

extreme values for wages and sales (0.1% top and bottom observations). After these exclusions, we obtained a sample of 106,549 workers that were displaced between 1994-96 due to firm closing.

The control group includes three sub-samples and was constructed in the following way. For each year prior to the displacement year we obtained a random sample of around 300,000 workers who were employed in firms that did not close.[10] These individuals were followed in the three years before displacement and in the years after displacement until 1998. The control group of the 1994 displacements includes a sample of workers employed in 1993 in firms that did not close in 1994. These workers were followed over the 1991-98 period if they remained with the same employer over that period. The control group of the 1995 displacements includes a sample of workers employed in 1994 in firms that did not close in 1995. These workers were followed over the 1992-98 period. Finally, the control group of the 1996 displacements includes a sample of workers employed in 1995 in firms that did not close in 1996. These workers were followed over the 1993-98 period if they remained with the same employer over that period. To be included in the sample a worker must also be present in the *QP* files in the year of displacement. In order to guarantee that the worker was employed with the same employer three years before separation, we control for worker's admission year in the firm.[11]

This sub-sample were also limited to full-time workers in the pre-displacement period aged between 18-59 in the year prior to displacement. After excluding those observations with missing values in the explanatory variables and the extreme values in wages and sales, we obtained a control group composed by 172,624 non-displaced workers.[12]

3.2 Some Basic Characteristics

Table 3.1 presents average real hourly earnings of displaced workers around the date of displacement.[13] The year of displacement is denoted as year zero. The years that precede the displacement are denoted as years minus three to minus one and the following years as years one through four. The pattern of average hourly earnings is shown for all these periods, separately for men and women. Comparable figures for non-displaced workers are also reported. In columns 5 (for men) and 10 (for women) the ratio of average hourly earnings of displaced and non-displaced workers is reported. Figure 3.1 displays the evolution of the ratio over the entire period of analysis.

[10] The sample was drawn according to a normal random number generator.

[11] In the year prior to displacement the worker must have at least two years of tenure with the employer.

[12] In Appendix A the basic characteristics of the samples of displaced and non-displaced workers are presented separately for both men (Tables A.3.1 and A.3.2) and women (Tables A.3.3 and A.3.4).

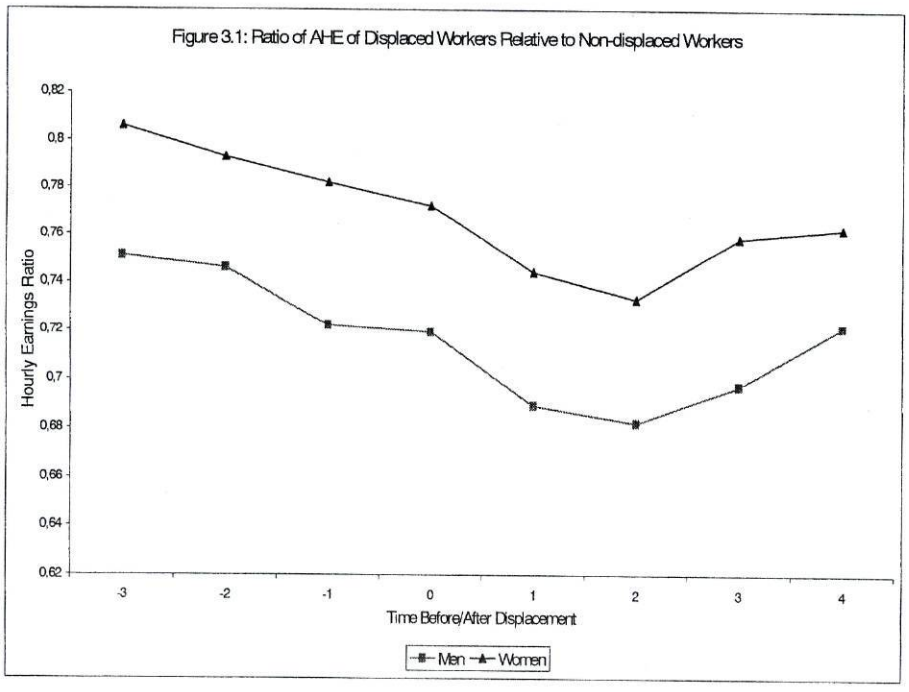
[13] Average hourly earnings are defined as the ratio between total regular payroll and the total number of usual hours worked.

As can be shown in Table 3.1, displaced workers earn, on average, lower wages than continuously employed workers. This gap is more pronounced for male workers. Moreover, and as is made clear in Figure 3.1, the earnings differential between displaced and non-displaced workers tends to increase with the approximation of the displacement year and deepens until two years after displacement, when a pattern of recovery appears to emerge. This analysis seems to reveal that a displacement event tends to deepen the earnings gap between displaced and non-displaced workers. However, it can always be argued that displaced workers would have had a slow earnings growth even without being displaced. A regression analysis will be required in order to control for observed individual and firm characteristics and to be able to determine the wage growth that displaced workers would have experienced if they had not been displaced.

Table 3.1: Average Real Hourly Earnings Before and After Displacement

	Men					Women				
	Displaced		Non-displaced		Ratio	Displaced		Non-displaced		Ratio
Years Before/After	AHE	Obs.	AHE	Obs.	(1)/(3)	AHE	Obs.	AHE	Obs.	(6)/(8)
Displacement	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
3 years before	421.1	38850	561.0	75954	0.751	322.2	28845	399.8	45459	0.806
2 years before	435.3	37441	583.7	89110	0.746	332.9	28829	419.9	55330	0.793
1 year before	430.8	31276	597.0	105506	0.722	337.7	25411	431.7	67118	0.782
Displacement year	448.3	6442	623.2	105506	0.719	348.3	4433	451.0	67118	0.772
1 year after	445.1	10302	646.3	88994	0.689	346.1	7660	465.1	55325	0.744
2 years after	460.7	12077	675.6	83548	0.682	356.1	8897	485.6	51790	0.733
3 years after	475.0	9422	682.0	49626	0.697	368.2	7168	485.8	29924	0.758
4 years after	500.7	5679	694.3	23371	0.721	379.4	4510	497.8	13677	0.762

Notes: (i) AHE is average real hourly earnings [PTE (escudo): 1 EUR≡200.482 PTE]; CPI deflator (base=1991);
(ii) Obs. denotes number of observations.



Finally, some additional considerations should be made concerning the composition of the sample of displaced workers. First, it should be observed that of the total number of displaced workers, 47% left the firm one or two years prior to the displacement year (see Table 3.2). Of the total number of workers that left the firm one year before displacement, 30% were reemployed. This rate is only 22% for workers who left their firms two years before displacement. These results seem to suggest that early leavers have lower reemployment rates. It would be interesting to know if these workers left the firm voluntarily or because they were laid off. In this data set, however, it is not possible to know the causes of workers' separations.

Second, of the total number of reemployments 58%, experience a positive spell of nonemployment and only 4% two or more spells (see Table 3.3). Furthermore, and according to Table 3.3, 35% of the reemployed individuals changed industry.[14] After being reemployed, 22% of the reemployed individuals changed job again (19% one time and 3% two or more times).

Third and finally, it should be noted that of the total number of workers that were displaced between 1994 and 1996, only 33% were reemployed in the following period. Of this total, 44% were immediately reemployed in the dis-

[14] Here industry is defined at the one-digit level. According to the Portuguese Classification of Economic Activities (CAE), at the one-digit level there are nine sectors. We decided to define the industry at the one-digit level in order to avoid measurement errors, since in 1995 the CAE codes were altered.

placement year, 26% one year after, 17% two years after, 9% three years after and the remaining 4% only four years after the displacement occurred (see Table 3.4). This low rate of reemployment may be justified, on the one hand, by the existence of a high replacement rate of the unemployment benefit and, on the other hand, by the long period of duration of the unemployment benefit.[15] Thus, it is possible that the current profile of the unemployment benefit system, which brings together a high replacement rate and a perspective of a long maximum period for the granting of the benefit, may contribute to reducing the transitions from unemployment to employment.

Nonetheless, these figures point to a potential problem of selectivity that may arise when we analyze the earnings losses of only those displaced workers that reenter employment. In fact, for an individual's wage to be observed in the post-displacement period he/she must be employed and registered in the *QP* files at least one time over that period. This requirement may give rise to non-randomness in the sample of wage observations, and if displaced reemployed workers have (un)observed characteristics that are systematically different for those who stay out of employment, this sample selection may bias the earnings losses estimates. In order to check that the earnings losses estimates are not biased due to sample selection, in the next Section a selectivity test for the sample of displaced workers will be presented.

[15] In Portugal the unemployment benefit represents approximately 65 per cent of the wages that an individual expects to obtain in the labor market. The duration of this benefit varies from a maximum of 12 months for individuals under 25 years of age to a maximum of 36 months for those aged 55 or over.

See the Annex for a detailed description of the unemployment protection scheme in Portugal.

Table 3.2: Early-leavers - Timing and Frequency

Displaced Workers [N=106549]			
Early Leavers [N=49862 (47%)]			
2 Years Before [N=23304 (47%)]		1 Year Before [N=26558 (53%)]	
Reemp.	Non Reemp.	Reemp.	Non Reemp.
5235 (22%)	18069 (78%)	7942 (30%)	18616 (70%)

Table 3.3: Reemployments - Some Basic Characteristics

Displaced Workers [N=106549]				
Reemployed [N=34825 (33%)]				
Out of Employment		Changed Job After Reemp.		Changed
1 spell	2 or more	1 time	2 or more	Industry
20311 (58%)	1221 (4%)	6641 (19%)	1016 (3%)	12155 (35%)

Table 3.4: Reemployments by Years since Displacement

Displaced Workers [N=106549]				
Reemployed [N=34825 (33%)]				
Disp. Year	1 Year After	2 Years After	3 Years After	4 Years After
15623 (44%)	8896 (26%)	6057 (17%)	3257 (9%)	1352 (4%)

4 Empirical Model and Results

4.1 The Econometric Model

As mentioned before, the main goal in this study is to compare the earnings paths of displaced and continuously employed workers in the period before and after the displacement has occurred. The empirical model is drawn from Jacobson *et al.* (1993) and Margolis (1999). An equation capturing the difference in earnings across the displaced and comparable non-displaced workers in the sample can be written as:

$$\ln W_{it} = \beta_1 X_{it} + \beta_2 Z_{ijt} + \sum_{k=-3}^{4} D_{it}^{k}\delta_k + \varphi \sum_{\iota=1}^{4} OE_{i\iota} + \gamma_t + \varepsilon_{it} \qquad (3.1),$$

where $\ln W_{it}$ is the natural log of average real hourly earnings for individual i at period t. Average hourly earnings are defined as the ratio between total regular payroll and the total number of usual hours worked.[16] Average hourly earnings were deflated using the Consumer Price Index (CPI; base=1991). X_{it} is a vector of individual characteristics such as education and age. Z_{ijt} includes a set of characteristics from the firm in which the individual works such as size and sales per worker.

D_{it}^{k} is a dummy variable that takes the value one if at time t worker i is k years after separation or -k years before separation. The δ_k parameters reflect the difference in earnings k years before or after separation between displaced workers and the corresponding reference group. Hereinafter, we will refer to these dummies as displacement dummies. The variable $OE_{i\iota}$ takes the value one in the after-separation period for those individuals who experience at least one spell of nonemployment (zero otherwise). The parameter φ captures a negative permanent effect on earnings for those individuals who experienced a nonemployment spell.

γ_t is a set of time dummies that control for year-specific effects. ε_{it} is a disturbance term with zero mean and constant variance.

Equation (3.1) will be estimated by ordinary least squares (OLS) separately for the samples of male and female workers. The validity of this estimation method relies on the assumption that the included independent variables control for all the individual heterogeneity that could be correlated with the occurrence of a displacement. In this study the fact that job displacement occurs for exogenous reasons (firm closing) makes this assumption reasonable. If unobserved permanent individual heterogeneity affects the probability of displacement and is correlated with wage determinants, OLS estimates of equation (3.1) will be biased.

If unobserved individual heterogeneity is of concern, the model could be estimated by fixed effects. Even though the fixed effects model allows us to control for unobserved heterogeneity, the inexistence of time variability does not allow us to identify the coefficients of the time-invariant variables, invalidating the

[16] Total regular payroll includes base wages, seniority payments and regular benefits.

decomposition of the effects of workers and firm characteristics in the difference in earnings between displaced and continuously employed workers.

4.2 Empirical Results

4.2.1 Pre- and Post-displacement Earnings Losses - OLS Regressions

In Tables 3.5 and 3.6 the OLS regression results of model (3.1) are presented for men and women, respectively. Column 1 of Tables 3.5 and 3.6 provides results for a parsimonious specification in which average real hourly earnings depend on the displacement dummies ($Disp_{it}$) and a set of individual characteristics that do not change with the job, such as age (and its square) and education. Education is defined as the number of years of schooling completed. Two dummy variables were added to the model: one that takes the value one if the worker has a part-time job in the post-displacement period (zero otherwise), and the other that takes the value one for displaced workers who left the firm one or two years before closing (zero otherwise). A set of time dummies is also included in order to control for macroeconomic conditions.

As can be shown in column 1 of Table 3.5, men that will be displaced in year zero earn, three years before separation, 12.8% less than their non-displaced counterparts, conditional on age, education, and macroeconomic conditions.[17] In these same conditions women earn 9.3% less (see column 1 of Table 3.6). This gap increases for both male and female with the approximation of the displacement year. Two years after the displacement year, Portuguese men earn 22.3% less than their continuously employed counterparts, and women earn 17.8% less. Only in the latter years does a slower recovery in earnings seems to emerge. In any event, three years after displacement men's earnings differential has risen by around 10 percentage points (p.p.) and women's differential by around 8 p.p., when compared to the earnings differential three years before displacement.[18] This pattern is, in general, consistent with Figure 3.1. However, the introduction of a control for worker characteristics considerably decreases the earnings differential, suggesting that differences in worker characteristics between the two groups explain a sizable part of the wage gap between displaced and non-displaced workers.

Early-leavers earn only 1.1-1.3% less than other similar displaced workers. Even though early-leavers represent almost half of the sample of displaced workers, the pattern of earnings losses for these workers does not differ substantially

[17] To calculate the percentage change in y induced by $\triangle x$, i.e. $\frac{y_{t+1}-y_t}{y_t}$, the exact formula is $\frac{y_{t+1}-y_t}{y_t} = \exp(\delta) - 1$.

[18] In Table B.3.1 of Appendix B this same parsimonious specification was estimated with individual-specific fixed effects. It shoud be noted that the fixed effects estimates of the coefficients of the displacement dummies do not have a straightforward interpretation in terms of earnings changes of displaced workers relative to non-displaced workers, since the coefficients represent within-individual earnings changes. Nevertheless, the fixed effects estimates exhibit the same qualitative result - four years after displacement, average hourly earnings of displaced workers are somewhat below their pre-displacement levels.

from the pattern of earnings losses of those workers who remain in the firm until the closing down.[19]

Differences in earnings between displaced and similar non-displaced workers may be explained by differences in employer characteristics where the individuals work. It is well established that closing firms have some characteristics (such as size and age) that are systematically different from those that remain in activity. It is also well known that differences in employer's performance may affect wages. In order to analyze to what extent the earnings gap between displaced and non-displaced workers is determined by differences in employer characteristics, a set of controls for employers observing heterogeneity was included in column 2 of Tables 3.5 and 3.6. The first variable is firm size, which is measured as the natural log of total employment in the firm. Sales per worker is defined as the ratio of annual real sales and total employment (in logs). Finally, eight industry (one-digit level) and six regional (NUTs II) dummies were added to the model.

The effects of size and sales per worker on average hourly earnings are positive and strongly significant. The results reveal that a large part of the relative annual earnings differential may be explained by differences in employers observed characteristics. Three years before separation the earnings gap between displaced workers and the reference group is negligible and not statistically different from zero (+0.4% for men and +0.3% for women). Nonetheless, and after controlling for firms' characteristics, it is still possible to observe a very similar pattern on earnings evolution over the entire period of analysis. A pre-displacement dip in earnings is observed, followed by a drop in earnings in the displacement year and a slower recovery four years after separation (see column 2 of Tables 3.5 and 3.6). Not surprisingly, three years after displacement the relative earnings differential has risen by around 11 p.p. for men and 9 p.p. for women (almost the same values as obtained in specification 1).

These results point to the importance of controlling for observed employers' characteristics when analyzing the effects of a displacement event on earnings. In fact, firms' characteristics such as size and sales per worker have a significant impact on earnings, and not controlling for them may confound the evaluation of the earnings losses. In the next Section we shall decompose the difference in earnings between displaced and non-displaced workers according to firm characteristics.

[19] The same regressions presented in Tables 3.5 and 3.6 were re-estimated for a subsample of workers in which early-leavers were excluded. The results are quite similar to those obtained for the overall sample, and are not recapitulated here (available upon request).

Table 3.5: Ordinary Least Squares Regressions - Men
Dependent variable: log of Average Real Hourly Earnings (N=773104)

	(1)		(2)		(3)		(4)	
Independent Variables	Coef.	t-ratio	Coef.	t-ratio	Coef.	t-ratio	Coef.	t-ratio
$Disp_{-3}$	-0.137	(-49.9)	0.004*	(1.9)	-0.003*	(-1.3)	0.004*	(1.7)
$Disp_{-2}$	-0.159	(-62.7)	-0.013	(-5.8)	-0.018	(-8.4)	-0.012	(-5.5)
$Disp_{-1}$	-0.199	(-80.9)	-0.033	(-15.8)	-0.033	(-15.7)	-0.030	(-14.3)
$Disp_0$	-0.207	(-39.0)	-0.079	(-17.4)	-0.043	(-9.2)	0.008*	(1.6)
$Disp_1$	-0.239	(-56.4)	-0.093	(-25.7)	-0.046	(-11.7)	0.006*	(1.5)
$Disp_2$	-0.252	(-63.8)	-0.114	(-33.7)	-0.055	(-13.8)	-0.0008*	(-0.2)
$Disp_3$	-0.257	(-57.6)	-0.118	(-31.0)	-0.054	(-11.9)	0.0009*	(0.2)
$Disp_4$	-0.254	(-44.4)	-0.116	(-23.9)	-0.048	(-8.6)	0.008*	(1.4)
Age	0.061	(187.9)	0.045	(163.2)	0.045	(163.2)	0.043	(148.1)
Age Squared[a]	-0.055	(-134.6)	-0.042	(-121.2)	-0.042	(-121.2)	-0.043	(-116.3)
Education	0.098	(639.0)	0.068	(460.5)	0.068	(460.7)	0.069	(469.6)
Part-time Job	0.096	(26.6)	0.090	(29.2)	0.093	(29.9)	0.093	(30.2)
Early-leaver	-0.013	(-5.5)	-0.028	(-13.4)	-0.014	(-6.4)	-0.011	(-5.0)
Size			0.064	(259.0)	0.064	(259.3)	0.059	(231.4)
Sales per Worker			0.100	(234.9)	0.100	(234.6)	0.101	(237.2)
OE (Out of Employment)					-0.069	(-19.0)	-0.062	(-17.2)
JC (Job Changes)					-0.015	(-3.0)	-0.013	(-2.6)
CIND (Changed Industry)					-0.062	(-17.5)	-0.066	(-18.7)
Tenure							0.007	(37.6)
Tenure Squared[a]							-0.003	(-5.3)
Constant	4.115	(651.8)	3.532	(554.7)	3.533	(555.2)	3.566	(556.7)
$\overline{R}^2$	0.44		0.60		0.60		0.60	

Notes: (i) all specifications include a set of time dummies;

(ii) specifications (2), (3) and (4) include a set of industry and regional dummies;

(iii) [a] variables divided by 100; (iv) all estimates are significant at 1%, except those with an *.

Table 3.6: Ordinary Least Squares Regressions - Women
Dependent variable: log of Average Real Hourly Earnings (N=501494)

	(1)		(2)		(3)		(4)	
Independent Variables	Coef.	t-ratio	Coef.	t-ratio	Coef.	t-ratio	Coef.	t-ratio
$Disp_{-3}$	-0.098	(-34.6)	0.003*	(1.3)	-0.003*	(-1.4)	0.007	(3.1)
$Disp_{-2}$	-0.109	(-42.5)	-0.003*	(-1.3)	-0.008	(-3.4)	0.002*	(0.8)
$Disp_{-1}$	-0.133	(-54.5)	-0.013	(-6.1)	-0.012	(-5.8)	-0.008	(-4.1)
$Disp_0$	-0.153	(-27.2)	-0.052	(-10.8)	-0.010*	(-2.0)	0.059	(11.9)
$Disp_1$	-0.178	(-41.0)	-0.065	(-17.5)	-0.014	(-3.4)	0.054	(13.3)
$Disp_2$	-0.196	(-48.1)	-0.091	(-26.6)	-0.030	(-7.6)	0.038	(9.4)
$Disp_3$	-0.194	(-42.7)	-0.096	(-25.1)	-0.026	(-5.6)	0.042	(9.1)
$Disp_4$	-0.188	(-32.8)	-0.090	(-18.6)	-0.018	(-3.2)	0.048	(8.6)
Age	0.037	(101.3)	0.026	(83.3)	0.026	(83.4)	0.019	(58.9)
Age Squared[a]	-0.028	(-58.4)	-0.020	(-48.6)	-0.020	(-48.8)	-0.014	(-33.7)
Education	0.096	(553.7)	0.068	(390.8)	0.067	(391.0)	0.068	(398.8)
Part-time Job	0.139	(40.8)	0.096	(33.2)	0.101	(35.1)	0.103	(36.1)
Early-leaver	-0.011	(-4.2)	-0.025	(-11.6)	-0.011	(-5.1)	-0.008	(-3.7)
Size			0.059	(229.5)	0.059	(230.7)	0.054	(209.9)
Sales per Worker			0.087	(184.7)	0.086	(184.1)	0.084	(179.2)
OE (Out of Employment)					-0.051	(-13.8)	-0.036	(-9.8)
JC (Job Changes)					-0.019	(-3.6)	-0.011*	(-2.2)
CIND (Changed Industry)					-0.125	(-32.2)	-0.127	(-33.0)
Tenure							0.013	(62.0)
Tenure Squared[a]							-0.025	(-36.8)
Constant	4.363	(663.2)	3.837	(581.9)	3.838	(582.8)	3.950	(593.1)
$\overline{R}^2$	0.45		0.61		0.61		0.62	

Notes: see notes to table 3.5.

In column 3 of Tables 3.5 and 3.6 controls for time spent out of employment and for post-displacement mobility were added to specification 2. The variable *OE* takes on the value one in all years following the displacement if the individual experienced one or more spells of nonemployment. A worker experiences a positive spell of nonemployment if he (she) is employed in year t and is not immediately reemployed in year $t+1$.[20]

Since our sample includes displaced workers that may change job several times after being reemployed, a variable that controls for the effects of changing a job on subsequent wages was also included in equation (3.1). An individual experiences a job change if the identification number of his recent employer is different from the identification number of his previous employer. The variable *JC* takes the value one in all years after a job change has occurred, if the individual changed job again after being reemployed. Two categories were defined: 1 or more job changes and no job change (the omitted category). Finally, a dummy variable for industry changers was included in column 3. The variable *CIND* takes the value one in the post-displacement period for reemployed displaced workers who changed industry (at the one-digit level).

The results reported in column 3 of Tables 3.5 and 3.6 show that workers who experienced one or more spells of nonemployment are those most affected by displacement. In fact, once a control for time spent out of employment is included, the displacement dummies reflect the difference in average hourly earnings between the control group and those displaced workers who do not experience any spell of nonemployment. For the latter, the annual earnings gap three years after displacement increased by 5 and 2 p.p. for men and women, respectively. Male workers who experienced at least one spell of nonemployment have an additional penalty on hourly earnings of 6.9 p.p. and female workers of 5.1 p.p.

This result seems to suggest that earnings losses for displaced workers in Portugal reflect a major difference between those who find a new job quickly and those who do not. This empirical evidence is in line with that obtained by Bender *et al.* (1999) and Margolis (1999) for displaced workers in Germany and France.

Furthermore, the results reveal that workers reemployed in the same one-digit industry suffered lower post-displacement earnings losses. In fact, displaced male workers who changed industry earn 6.2% less and displaced women 12.5% less. These results are in line with those obtained by Addison and Portugal (1989), Jacobson *et al.* (1993), Carrington (1993) and Stevens (1997), who showed that industry changers in the U.S experience larger earnings losses. Finally, male workers who after being reemployed changed job again, have an earnings disadvantage of around 1.5% and female workers of 1.9%.

Post-displacement earnings losses may stem from the loss of accumulated firm-specific human capital and/or a "good" job match. In column 4 of Tables 3.5 and 3.6 a control for worker's tenure was added to specification 3. Tenure

[20] Note that we are also controlling in the regressions for subsequent separations after the first displacement through firm closing.

is measured as the total number of years with the employer. The inclusion of a control for the returns to accumulated tenure will enable us to determine the extent to which post-displacement earnings losses result from the loss of firm-specific capital and/or a high quality match.

The results reveal that tenure has an important impact on the magnitude of the relative annual earnings differential. As expected, the post-separation earnings differentials are much smaller (by around 5 to 7 p.p.) when a control for job tenure is included in the model. According to column 4, once losses of firm tenure are accounted for, no negative annual earnings differential is observed over the post-displacement period for either male or female workers. These results suggest that the loss of match-specific characteristics has a significant impact on earnings, deepening the average hourly earnings gap between displaced and their similar non-displaced counterparts.

To sum up, we can decompose the earnings losses of displaced workers relative to non-displaced workers into three components: (i) the loss associated with nonemployment duration; (ii) the loss that stems from the loss of tenure in the job; and (iii) the loss related with industry changing. Overall, the results for the male sample reveal that the increase in the earnings gap of 11 p.p. is mainly due to the loss of tenure in the job and to joblessness. Indeed, joblessness and tenure account equally (42% each) for the deepening in the earnings gap. Changing industry explains only 16% of the increase in the earnings gap. For female workers, the increase in the earnings gap of 9 p.p. is mainly due to the loss of accumulated returns to tenure (50%), joblessness accounts for 36% of that increase and changing industry accounts only for 14%.

The four regressions presented in Tables 3.5 and 3.6 were re-estimated for a sub-sample of workers in which minimum wage earners were excluded and for a sub-sample of workers in which displaced workers that changed job again after being re-employed were dropped (separately for men and women). These results are quite similar to the ones obtained in tables 3.5 and 3.6 and are not recapitulated here (available upon request). Finally, in Appendix C we present the results for the selectivity test that was carried out for the sample of displaced workers.[21]

4.2.2 Pre- and Post-displacement Earnings Losses - Quantile Regressions

In this Section we use quantile regression techniques in order to examine if displacement effects on wages vary along the wage distribution. Unlike the conditional mean regression, the quantile regression analysis allows us to determine the effect of each of the covariates along the whole distribution of hourly wages

[21] The results reveal that even though the selectivity term is significantly different from zero, it does not affect appreciably the parameter estimates or t-statistics of the independent variables and, in particular, the displacement dummy variables.

An alternative route to correct for selectivity would be to employ the inverse probability weighted regression method suggested by Wooldridge (2002).

and to examine to what extent the marginal effect of displacement on wages depends on the point of the conditional wage distribution where the individual is located.

Table 3.7 presents the quantile regressions estimates for male workers according to specifications 2 and 4 of Table 3.5. Table 3.8 presents the same estimates for female workers. The selected quantiles are 10-th, 50-th and 90-th. The estimated coefficients measure the impact of each covariate on wages for each specific quantile.

According to column 1 of Tables 3.7 and 3.8, even though the effect of each displacement dummy differs across quantiles, the earnings gap between displaced and similar non-displaced workers increased over the period by almost the same amount at the top, median and bottom of the wage distribution. Indeed, three years after displacement, and compared to the earnings gap three years before displacement, the annual earnings differential between displaced male workers and their counterparts has increased by 11 p.p. at the first and ninth decile and by 10 p.p. at the median. For female workers, the same earnings gap increased by 8 p.p. at the first decile, by 7 p.p. at the median and by 9 p.p. at the top decile. These results seem to suggest that the magnitude of the full effect of displacement on earnings is fairly similar across the whole wage distribution. Yet, once controls for job tenure and for time spent out of employment are included in the model, some differences across quantiles seem to emerge, namely for male workers.

According to column 2 of Table 3.7, once losses of accumulated returns to tenure and loss of sector-specific factors are accounted for, no negative post-displacement earnings differential appears to exist between displaced and similar non-displaced male workers at the 10-th and 50-th quantile. A dip in earnings is observed only in the pre-displacement period. At the 90-th decile, however, male workers seem to suffer post-displacement earnings losses even after controlling for job tenure and industry changing.[22] This might be justified by the fact that the returns to tenure decrease for higher quantiles, suggesting that tenure is more valued at relatively low paid jobs. On the contrary, the returns to education are higher for the top quantiles.[23]

In sum, the results for the sample of male workers seem to suggest that even though the magnitude of the global effect of displacement on wages is quite similar across quantiles, the sources of these losses seem to differ along the wage distribution. While for low paid male workers, tenure within the firm and industry changing explain most of the earnings losses experienced by these workers, for high paid male workers there are other (unobserved) factors beyond tenure/industry changing that account for a non-negligible amount of these losses. Changing industry imposes a larger penalty on male earnings at the bottom than at the top quantiles. For low and high paid male workers the effects of being out of employment are quite similar, even though weaker for the

[22] According to column 4 of Table 3.7, the coefficients of the post-displacement dummies, although negative, are not statistically different from zero at the 1% level of significance.

[23] See Machado and Mata (1998) for a detailed description of the conditional wage distribution for Portugal and its evolution from 1982 to 1994.

former.

For the sample of female workers, as before, once the returns to accumulated tenure are accounted for, a positive earnings differential between displaced (immediately reemployed) and non-displaced workers is observed in the post-displacement period. In this case, and contrary to the sample of male workers, the displacement effects are quite similar across all the selected quantiles (see Table 3.8).

Overall, the regression quantile results reinforce the OLS results obtained in the previous Section, showing that the full effect of displacement on earnings is quite similar across quantiles. Low and high paid workers seem to be equally affected, in terms of magnitude and persistence, by job displacement. The only exception concerns the different sources of wage adjustment. For female and low paid male workers who find a new job quickly, the main source of earnings adjustment stems from the loss of accumulated returns to tenure. For high paid male workers, and since tenure on the job is less valued than for low paid workers, there are other unobserved sources of earnings losses. One possiblity is that these latter losses may stem from the loss of search rents earned on the job lost. This type of explanation was advanced by Lefranc (2002) using a job search approach. The basic idea of the model is that wages will fall after displacement because before displacement workers have, on average, access to better jobs than when exiting unemployment due to the process of on-the-job-search that leads to the accumulation of "search rents". According to his model, after controlling for the accumulation of job-specific human capital, wage losses can be used to assess the level of reservation wages. Indeed, as workers increase their reservation wages, the range of acceptable wage offers shrinks and, consequently, the range for profitable job-to-job mobility will decline. Thus, loss of search rents is negatively related to the level of the reservation wage. Since in Portugal the unemployment benefit is considerably high, low paid workers may have high reservation wages relative to high paid workers, which leads to high unemployment duration for the former, limiting the range of profitable job-to-job mobility and consequently the losses of search rents after displacement.[24] On the contrary, for high paid workers, and due to the existence of relatively low reservation wages, the loss of search rents may be considerably higher.

[24] The amount of benefit paid represents around 65% of the average monthly work pay during a 6-month period prior to the second month before becoming unemployed. This amount can neither be less than the legal minimum wage nor exceed three times its value.

Table 3.7: Quantile Regressions - Men

Dependent variable: log of Average Real Hourly Earnings (N=773104)

	(1)						(2)					
	Q=0.1		Q=0.5		Q=0.9		Q=0.1		Q=0.5		Q=0.9	
Independent Variables	Coef.	t-ratio	Coef.	t-ratio	Coef.	t-ratio	Coef.	t-ratio	Coef.	t-ratio	Coef.	t-ratio
$Disp_{-3}$	-0.012*	(-1.4)	0.002*	(0.9)	0.014*	(1.5)	-0.010*	(-1.2)	0.003*	(1.6)	0.011*	(1.2)
$Disp_{-2}$	-0.025	(-3.2)	-0.015	(-9.0)	-0.004*	(-0.4)	-0.022	(-2.8)	-0.013	(-7.9)	-0.004*	(-0.5)
$Disp_{-1}$	-0.042	(-5.5)	-0.035	(-22.3)	-0.017*	(-2.0)	-0.038	(-5.0)	-0.032	(-20.5)	-0.017*	(-2.0)
$Disp_0$	-0.089	(-5.5)	-0.076	(-22.8)	-0.062	(-3.4)	0.033*	(1.9)	0.016	(4.4)	-0.014*	(-0.7)
$Disp_1$	-0.103	(-7.9)	-0.082	(-30.9)	-0.070	(-4.8)	0.036*	(2.5)	0.018	(6.0)	-0.020*	(-1.2)
$Disp_2$	-0.115	(-9.5)	-0.100	(-40.1)	-0.106	(-7.8)	0.040	(2.8)	0.011	(3.8)	-0.034*	(-2.0)
$Disp_3$	-0.124	(-9.1)	-0.102	(-36.3)	-0.103	(-6.7)	0.035*	(2.2)	0.014	(4.2)	-0.029*	(-1.5)
$Disp_4$	-0.120	(-6.9)	-0.104	(-28.9)	-0.106	(-5.4)	0.040*	(2.0)	0.016	(4.0)	-0.030*	(-1.3)
Age	0.038	(37.7)	0.044	(213.1)	0.053	(46.9)	0.032	(31.0)	0.041	(189.6)	0.054	(45.5)
Age Squared[a]	-0.037	(-29.8)	-0.042	(-161.3)	-0.047	(-33.4)	-0.034	(-26.1)	-0.041	(-152.4)	-0.051	(34.3)
Education	0.042	(79.6)	0.063	(581.6)	0.091	(153.4)	0.043	(82.4)	0.064	(594.1)	0.093	(155.6)
Part-time Job	-0.0003*	(-0.0)	0.044	(19.5)	0.251	(20.2)	0.005*	(0.4)	0.045	(20.1)	0.260	(21.0)
Early-leaver	-0.050	(-6.8)	-0.035	(-23.2)	0.008*	(1.0)	-0.025	(-3.4)	-0.021	(-13.1)	0.021*	(2.5)
Size	0.059	(67.2)	0.063	(347.6)	0.061	(60.7)	0.053	(59.4)	0.057	(306.9)	0.056	(55.0)
Sales per Worker	0.107	(69.9)	0.104	(329.8)	0.092	(53.7)	0.107	(71.1)	0.105	(335.9)	0.093	(54.1)
OE (Out of Employment)							-0.051	(-4.0)	-0.050	(-19.1)	-0.065	(-4.4)
JC (Job Changes)							-0.037*	(-2.2)	-0.015	(-4.3)	0.009*	(0.4)
CIND (Changed Industry)							-0.090	(-7.2)	-0.059	(-22.8)	-0.044	(-3.1)
Tenure							0.012	(19.5)	0.008	(60.5)	0.001*	(1.6)
Tenure Squared[a]							-0.016	(-8.7)	-0.005	(-13.4)	0.011	(5.0)
Constant	3.462	(152.0)	3.545	(754.9)	3.708	(144.5)	3.537	(155.9)	3.590	(762.2)	3.705	(143.4)
$\overline{R}^2$	0.58		0.59		0.58		0.58		0.60		0.59	

Notes: (i) all specifications include a set of industry, regional and time dummies; (ii) [a] variables divided by 100;

(iii) all estimates are significant at 1%, except those with an *.

Table 3.8: Quantile Regressions - Women

Dependent variable: log of Average Real Hourly Earnings (N=501494)

	(1)						(2)					
	Q=0.1		Q=0.5		Q=0.9		Q=0.1		Q=0.5		Q=0.9	
Independent Variables	Coef.	t-ratio	Coef.	t-ratio	Coef.	t-ratio	Coef.	t-ratio	Coef.	t-ratio	Coef.	t-ratio
$Disp_{-3}$	-0.003*	(-0.3)	0.005	(2.8)	0.012*	(1.3)	0.001*	(0.2)	0.007	(4.2)	0.016*	(1.7)
$Disp_{-2}$	0.003*	(0.5)	-0.004	(-2.9)	0.002*	(0.3)	0.009*	(1.2)	-0.002*	(-1.1)	0.007*	(0.8)
$Disp_{-1}$	-0.014*	(-2.0)	-0.014	(-9.5)	0.002*	(0.3)	-0.010*	(-1.5)	-0.013	(-9.0)	0.009*	(1.1)
$Disp_0$	-0.057	(-3.6)	-0.038	(-11.6)	-0.023*	(-1.2)	0.043	(2.7)	0.066	(19.4)	0.049	(2.6)
$Disp_1$	-0.066	(-5.5)	-0.045	(-17.7)	-0.034*	(-2.4)	0.044	(3.3)	0.066	(23.7)	0.051	(3.3)
$Disp_2$	-0.083	(-7.3)	-0.066	(-27.8)	-0.072	(-5.4)	0.032*	(2.4)	0.052	(18.7)	0.023*	(1.5)
$Disp_3$	-0.091	(-7.2)	-0.077	(-29.0)	-0.076	(-5.2)	0.029*	(1.9)	0.054	(16.9)	0.023*	(1.3)
$Disp_4$	-0.082	(-5.2)	-0.062	(-18.6)	-0.073	(-3.9)	0.038*	(2.1)	0.063	(16.4)	0.030*	(1.4)
Age	0.022	(22.0)	0.021	(98.9)	0.024	(20.8)	0.016	(15.5)	0.015	(69.0)	0.021	(17.4)
Age Squared[a]	-0.023	(-17.3)	-0.017	(-58.7)	-0.013	(-8.2)	-0.017	(-12.2)	-0.012	(-41.0)	-0.013	(-8.1)
Education	0.032	(56.8)	0.060	(502.0)	0.092	(139.0)	0.033	(58.1)	0.061	(512.6)	0.094	(143.1)
Part-time Job	-0.011*	(-1.2)	0.039	(19.8)	0.256	(23.1)	-0.003*	(-0.4)	0.051	(25.6)	0.264	(24.1)
Early-leaver	-0.029	(-4.2)	-0.025	(-17.3)	-0.009*	(-1.1)	-0.017*	(-2.4)	-0.009	(-6.2)	0.005*	(0.6)
Size	0.050	(59.7)	0.054	(305.3)	0.055	(56.0)	0.049	(57.8)	0.050	(278.8)	0.050	(50.9)
Sales per Worker	0.061	(39.4)	0.082	(252.6)	0.090	(50.1)	0.060	(39.6)	0.079	(245.7)	0.087	(48.8)
OE (Out of Employment)							-0.029*	(-2.4)	-0.028	(-11.1)	-0.032*	(-2.3)
JC (Job Changes)							-0.024*	(-1.4)	-0.015	(-4.3)	0.016*	(0.8)
CIND (Changed Industry)							-0.124	(-9.8)	-0.143	(-54.0)	-0.109	(-7.4)
Tenure							0.012	(17.7)	0.013	(86.0)	0.009	(10.9)
Tenure Squared[a]							-0.031	(-14.0)	-0.023	(-50.7)	-0.006*	(-2.5)
Constant	4.131	(190.6)	4.012	(878.9)	3.977	(157.3)	4.192	(191.2)	4.110	(895.7)	4.046	(159.0)
$\overline{R}^2$	0.58		0.60		0.60		0.59		0.61		0.60	

Notes: see notes to table 3.7.

5 The Contribution of Worker and Firm Characteristics to Earnings Difference

In this Section we examine how worker and firm characteristics contribute to the difference in earnings between displaced and non-displaced workers before and after displacement. The average earnings differential between displaced and non-displaced workers before and after displacement can be written as:

$$\overline{\ln W}_{AD} - \overline{\ln W}_{BD} = (\overline{\ln W}^{D}_{AD} - \overline{\ln W}^{ND}_{AD}) - (\overline{\ln W}^{D}_{BD} - \overline{\ln W}^{ND}_{BD}) =$$

$$= \beta_1[(\overline{X}^{D}_{AD} - \overline{X}^{ND}_{AD}) - (\overline{X}^{D}_{BD} - \overline{X}^{ND}_{BD}) +$$

$$+\beta_2[(\overline{Z}^{D}_{AD} - \overline{Z}^{ND}_{AD}) - (\overline{Z}^{D}_{BD} - \overline{Z}^{ND}_{BD})] + [\delta_{AD} - \delta_{BD}] \qquad (3.2),$$

where $\overline{\ln W}^{D}_{AD}$ is the mean of the log of average real hourly earnings of displaced workers in the after-displacement period and $\overline{\ln W}^{ND}_{AD}$ the mean of the log of average real hourly earnings of non-displaced workers in the post-displacement period. $\overline{\ln W}^{D}_{BD}$ and $\overline{\ln W}^{ND}_{BD}$ represent the same means but in the pre-displacement period. $\overline{X}$ are the means of variables that measure worker characteristics such as education and age, and $\overline{Z}$ the means of variables that control for firm characteristics such as size and sales per worker. These means are defined for displaced (D) and non-displaced workers (ND) in the period after (AD) and before displacememt (BD). β_1 and β_2 are the coefficient estimates of the variables included in $\overline{X}$ and $\overline{Z}$, respectively. δ_{AD} and δ_{BD} are the coefficient estimates of the displacement dummies in the post- and pre-displacement period respectively.

In order to obtain the estimates of the β's and the δ's parameters the time periods around the date of displacement were divided into two groups: the pre-displacement period, which includes the three years that precede the displacement year; and the post-displacement period, which includes the displacement year and the following four years. The estimates of an earnings regression equation similar to the one presented in column 4 of Table 3.5 are presented in Table B.3.2 of Appendix B for both men and women. Since we are interested in examining the contribution of worker and firm characteristics for earnings difference before and after displacement, we chose a specification that includes all the controls for observed workers' and firms' heterogeneity.

In Tables 3.9 and 3.10 the contribution of each independent variable to the difference in earnings before and after displacement is presented for male and female workers, respectively.[25] The results report that the post-displacement

[25] The variable *early-leavers* was not included in Tables 3.9 and 3.10 because its coefficient is not statistically different from zero.

earnings differential between displaced and non-displaced workers increased by 3.5 log points for male workers and by 4.3 log points for female workers (see column 7). This increase in the earnings differential between displaced and non-displaced workers is related to the displacement event in two distinct ways. On one side, job displacement might affect the composition of the sample of displaced workers, since only 33% of the total number of displaced workers are reemployed in the post-displacement period. On the other side, and supposing that the sample of displaced workers remains unchanged concerning time-invariant characteristics, the displacement has an effect on variables that are associated with the job such as tenure, industry and firm size. The decomposition results show that both forces are at work in explaining the difference in differences in earnings.

The decomposition of the contribution of each variable to the difference in differences in earnings shows that the earnings differential between displaced and non-displaced workers due to differences in education became positive. Indeed, whereas in the pre-displacement period displaced workers are, on average, less educated than non-displaced workers, in the post-displacement period displaced workers have, on average, more education (6.7 years against 6.4 for men and 6.9 against 6.6 for women). In fact, the data reveal that around 23% of the displaced workers that were reemployed changed their educational level.[26] On average, the number of years of education of these workers rose from 5 to 7.9 years of completed schooling. This empirical evidence seems to suggest that displacement might have a positive impact on displaced workers, in the sense that it encourages the investment in human capital.

For both samples the average experience of displaced workers in the post-displacement period has declined leading to an increase in the earnings differential between displaced workers and the reference group. The negative contribution of age may result from the fact that older workers who were displaced left the labor force, taking advantage of early retirement schemes.

Besides the difference in earnings that stem from changes in the composition of the displaced workforce, there are other factors directly related to the loss of the job that cause post-displacement earnings losses. The results show that for both men and women the loss of tenure on the job explains why the earnings differential between displaced and non-displaced workers deepened in the post-displacement period. Changing industry after a displacement also contributes considerably to the increase in the earnings gap between displaced and non-displaced workers. Moreover, workers who experienced a spell of nonemployment have an additional and significant penalty on earnings. Once again, the decomposition of the contribution of each variable to the difference in earnings between displaced and non-displaced workers shows that the loss of tenure on the job, joblessness and industry changing are the main factors that explain the deepening in the earnings gap in the post-displacement period (see column 8).

[26] Of the total number of workers that obtained more education after the displacement, 45% changed from 4 to 6 years of completed schooling and 14% from 6 to 9 years. It should be noted that the mandatory schooling is 9 years.

Finally, the results reported in Tables 3.9 and 3.10 also reveal that there are differences in earnings even before displacement between displaced and non-displaced workers due to differences in firm characteristics in which the individuals work, and that tend to persist in the post-displacement period. In fact, displaced workers are employed in the pre-displacement period in firms with a smaller average size and with a lower average productivity (measured by sales per worker). In the post-displacement period, even though the earnings differential that stem from differences in average size is reduced, it remains considerably high (see column 7). Concerning sales per worker, the earnings differential in the post-displacement period is reduced for male workers and increased for female workers (see column 7).

This evidence reinforces the importance of controlling for employers' attributes when analyzing the displacement effects on earnings, since the pools of displaced and non-displaced workers seem to be very heterogenous concerning firm characteristics. In fact, there are systematic differences in earnings between displaced and non-displaced workers due to systematic differences in firm characteristics that exist even before displacement, and tend to persist after the displacement.

The empirical literature on plant closings has been showing that firms that will soon close are usually of small size and lower productivity, so that the differences in firm characteristics for the two groups of workers in the pre-displacement period are not surprising. In the post-displacement period the differences in the two groups are not so justifiable, unless it can be argued that workers displaced through firm closing carry with them a "displacement stigma" that leads firms (namely, high productivity firms) facing a queue of job applicants to prefer those individuals who did not experience a displacement and probably no spell of unemployment [Blanchard and Diamond (1994)].

In sum, the decomposition of the contribution of each variable to the difference in earnings corroborates the regressions results that the post-displacement increase in the relative earnings differential for displaced workers who are immediately reemployed mostly stems from the loss of tenure within the firm and, to a lesser extent, from the loss of sector-specific features. For workers who take a long time to find a new job the earnings differential is considerably higher. For these workers the additional losses may result from the depreciation of the stock of human capital (general or specific) that a spell of nonemployment imposes or from lower wage offers received as employers perceive that these workers are of lower productivity.

Table 3.9: The Contribution of Worker and Firm Characteristics to the Difference in Differences in Earnings
Men (N=773104)

	Displaced Before Disp.	Non-disp. Before Disp.	Displaced After Disp.	Non-disp. After Disp.	Differential Before Disp.	Differential After Disp.	Differences in Differences	
					$[(1)-(2)]*\beta$	$[(3)-(4)]*\beta$	(6)-(5)	|(6)-(5)|
Independent Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Age	35.526	37.968	36.518	40.858	-0.105	-0.187	-0.082	
Age Squared	1382.233	1547.232	1432.964	1775.028	0.071	0.147	0.076	0.006
Education	5.738	6.180	6.718	6.433	-0.030	0.020	0.050	0.050
Part-time Job	0.000	0.000	0.062	0.030	0.000	0.003	0.003	0.003
Size	3.321	4.892	3.521	4.849	-0.093	-0.078	0.015	0.015
Sales per Worker	8.446	8.923	8.574	8.982	-0.048	-0.041	0.007	0.007
OE	0.000	0.000	0.520	0.000	0.000	-0.034	-0.034	0.034
JC	0.000	0.000	0.146	0.000	0.000	-0.002	-0.002	0.002
CIND	0.000	0.000	0.359	0.000	0.000	-0.024	-0.024	0.024
Tenure	7.308	10.736	1.263	13.751	-0.024	-0.087	-0.063	
Tenure Squared	114.135	188.875	3.136	262.452	0.002	0.008	0.006	0.057
$\delta_{AD} - \delta_{BD}$							0.019	0.019
							-0.035	0.223

Table 3.10: The Contribution of Worker and Firm Characteristics to the Difference in Differences in Earnings
Women (N=501494)

	Displaced Before Disp.	Non-disp. Before Disp.	Displaced Before Disp.	Non-disp. Before Disp.	Differential Before Disp.	Differential After Disp.	Differences in Differences	
					$[(1)-(2)]*\beta$	$[(3)-(4)]*\beta$	(6)-(5)	|(6)-(5)|
Independent Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Age	32.760	34.212	33.403	37.249	-0.028	-0.073	-0.045	
Age Squared	1170.295	1262.721	1191.212	1479.469	0.013	0.040	0.027	0.018
Education	5.955	6.373	6.943	6.643	-0.028	0.020	0.048	0.048
Part-time Job	0.000	0.000	0.080	0.043	0.000	0.004	0.004	0.004
Size	3.417	4.583	3.612	4.591	-0.064	-0.050	0.010	0.010
Sales per Worker	8.103	8.527	8.086	8.579	-0.035	-0.041	-0.005	0.005
OE	0.000	0.000	0.517	0.000	0.000	-0.021	-0.021	0.021
JC	0.000	0.000	0.127	0.000	0.000	-0.002	-0.002	0.002
CIND	0.000	0.000	0.287	0.000	0.000	-0.036	-0.036	0.036
Tenure	6.930	9.212	1.290	12.306	-0.030	-0.143	-0.113	
Tenure Squared	101.001	147.367	3.210	215.256	0.011	0.051	0.040	0.073
$\delta_{AD} - \delta_{BD}$							0.050	0.050
							-0.043	0.267

6 Conclusion

In this study we have analyzed the long-term earnings costs of displacement for Portuguese displaced workers through firm closing. For this purpose a large representative data set that links employers and its employees was used.

The findings of this research suggest four primary conclusions. First, we find that the effects of displacement on earnings are considerable in magnitude and quite persistent. Three years after displacement the average hourly earnings differential of displaced workers relative to continuously employed workers has increased by around 9 and 11 percentage points for women and men respectively. A slight recovery in earnings is observed only in the last year. These estimates are near the lower bound for the U.S.

Second, the results also showed that workers who experienced some period of nonemployment are the most affected by job displacement. Indeed, workers who experienced a spell of nonemployment have an additional annual penalty on average hourly earnings of 4 to 6 percentage points.

Third, the empirical results revealed that the deepening in the earnings differential between displaced and non-displaced workers is in large part due to the loss of tenure within the firm and, to a lesser extent, to industry changing, suggesting that job displacement leads to the destruction of firm (industry)-specific features that positively influence wages. In fact, once losses of accumulated returns to tenure are accounted for, the post-displacement earnings gap between displaced workers who were immediately reemployed and non-displaced workers is nearly zero (or positive, in the case of female workers). Aside from the losses that stem from the loss of job tenure, workers who take a long time to find a new job suffer an additional penalty on earnings that is directly related with time spent out of employment. Nevertheless, we found that displacement might have a positive impact on displaced workers, since these workers seem to be encouraged to invest in human capital while unemployed. On average, by the end of the observation period, reemployed displaced workers increased their number of years of schooling completed by one year.

In sum, the results revealed that for male workers the loss of tenure on the job explains about 42% of the increase in the post-displacement earnings gap relative to non-displaced workers; joblessness accounts for 42% of that increase; and the remaining 16% is explained by the loss of sector-specific features. As for female workers, the loss of accumulated returns to tenure account for 50% of the increase in the earnings differential; joblessness accounts for 36%; and the loss of sector-specific features explains the remaining 14%.

Fourth and finally, the decomposition of the contribution of worker and firm characteristics to the difference in earnings showed that there are significant differences in earnings between displaced and continuously employed workers that are related to differences in firm characteristics between these two groups of individuals. In fact, the pre-displacement differences in earnings between displaced workers and the control group due to differences in firms characteristics such as size and sales per worker tend to persist in the post-displacement period. The regression results resemble this finding and point to the impor-

tance of controlling for employer characteristics when estimating the earnings differential between displaced and similar non-displaced workers, since a gap in earnings exists even before displacement due to systematic differences in firm characteristics for the two groups of workers. Ignoring them may confound the evaluation of the earnings losses.

APPENDIX A - Descriptive Statistics

Table A.3.1: Descriptive Statistics (1991-98)
Displaced Male Workers

	MEAN	STDV	MIN	MAX
Variables				
Age (in years)	35.8	10.7	16.0	64.0
Education (in years)	6.0	2.9	0.0	16.0
Tenure (in years)	5.6	7.1	0.0	50.0
Qualification Level (%)				
Manager and Highly Professional	0.020		0.0	1.0
Professional	0.019		0.0	1.0
Supervisors	0.053		0.0	1.0
Highly Skilled and Skilled	0.551		0.0	1.0
Semi-skilled and Unskilled	0.234		0.0	1.0
Apprentices	0.077		0.0	1.0
Non-defined	0.046		0.0	1.0
Average Real Hourly Earnings (in logs)	5.948	0.46	4.55	8.38
Firm Size (in logs)	3.379	1.69	0.00	9.99
Firm Sales per Worker (in logs)	8.483	1.19	0.25	15.51
Number of Observations	151489			
Number of Individuals	60827			

Table A.3.2: Descriptive Statistics (1991-98)
Non-displaced Male Workers

	MEAN	STDV	MIN	MAX
Variables				
Age (in years)	39.6	10.4	16.0	64.0
Education (in years)	6.3	3.2	0.0	16.0
Tenure (in years)	12.4	8.7	0.0	61.0
Qualification Level (%)				
Manager and Highly Professional	0.033		0.0	1.0
Professional	0.036		0.0	1.0
Supervisors	0.074		0.0	1.0
Highly Skilled and Skilled	0.591		0.0	1.0
Semi-skilled and Unskilled	0.203		0.0	1.0
Apprentices	0.032		0.0	1.0
Non-defined	0.032		0.0	1.0
Average Real Hourly Earnings (in logs)	6.257	0.55	4.92	8.45
Firm Size (in logs)	4.868	2.10	0.00	10.29
Firm Sales per Worker (in logs)	8.956	1.05	0.92	17.29
Number of Observations	621615			
Number of Individuals	105506			

Table A.3.3: Descriptive Statistics (1991-98)
Displaced Female Workers

	MEAN	STDV	MIN	MAX
Variables				
Age (in years)	32.9	9.5	16.0	64.0
Education (in years)	6.2	2.9	0.0	16.0
Tenure (in years)	5.3	6.7	0.0	45.0
Qualification Level (%)				
Manager and Highly Professional	0.009		0.0	1.0
Professional	0.009		0.0	1.0
Supervisors	0.021		0.0	1.0
Highly Skilled and Skilled	0.446		0.0	1.0
Semi-skilled and Unskilled	0.352		0.0	1.0
Apprentices	0.125		0.0	1.0
Non-defined	0.038		0.0	1.0
Average Real Hourly Earnings (in logs)	5.719	0.39	4.62	8.39
Firm Size (in logs)	3.472	1.74	0.00	9.87
Firm Sales per Worker (in logs)	8.098	1.17	-0.09	16.13
Number of Observations	115753			
Number of Individuals	45722			

Table A.3.4: Descriptive Statistics (1991-98)
Non-displaced Female Workers

	MEAN	STDV	MIN	MAX
Variables				
Age (in years)	35.9	9.7	16.0	64.0
Education (in years)	6.5	3.1	0.0	16.0
Tenure (in years)	11.0	8.1	0.0	62.0
Qualification Level (%)				
Manager and Highly Professional	0.014		0.0	1.0
Professional	0.018		0.0	1.0
Supervisors	0.028		0.0	1.0
Highly Skilled and Skilled	0.486		0.0	1.0
Semi-skilled and Unskilled	0.375		0.0	1.0
Apprentices	0.055		0.0	1.0
Non-defined	0.024		0.0	1.0
Average Real Hourly Earnings (in logs)	5.944	0.51	4.92	8.43
Firm Size (in logs)	4.587	1.98	0.00	10.29
Firm Sales per Worker (in logs)	8.557	1.08	0.92	15.49
Number of Observations	385741			
Number of Individuals	67118			

APPENDIX B - Regression Results: Men and Women

Table B.3.1: Fixed Effects Results
Dependent variable: log of Average Real Hourly Earnings

	Men		Women	
Independent Variables	Coef.	t-ratio	Coef.	t-ratio
$Disp_{-3}$	0.010*	(1.8)	0.007*	(1.1)
$Disp_{-2}$	-0.004*	(-0.7)	-0.003*	(-0.4)
$Disp_{-1}$	-0.029	(-5.1)	-0.017	(-2.6)
$Disp_0$	-0.043	(-6.8)	-0.029	(-4.0)
$Disp_1$	-0.045	(-7.4)	-0.031	(-4.6)
$Disp_2$	-0.050	(-8.3)	-0.036	(-5.3)
$Disp_3$	-0.059	(-9.6)	-0.039	(-5.7)
$Disp_4$	-0.068	(-10.6)	-0.037	(-5.2)
Age Squared[a]	-0.049	(-82.6)	-0.036	(-46.3)
Education	0.004	(8.6)	0.006	(12.0)
Part-time Job	0.126	(65.7)	0.125	(62.8)
Early-leaver	0.024*	(1.7)	-0.013*	(-0.8)
$\overline{R}^2$	0.88		0.86	
Number of Observations	773104		501494	

Notes: (i) a set of time dummies is included in the specification;
(ii) [a] variables divided by 100;
(iii) all estimates are significant at 1%, except those with an *.

Table B.3.2: OLS Results
Dependent variable: log of Average Real Hourly Earnings

	Men		Women	
Independent Variables	Coef.	t-ratio	Coef.	t-ratio
$Disp_{BD}$	-0.016	(-10.8)	-0.002*	(-1.1)
$Disp_{AD}$	0.003*	(1.0)	0.048	(17.0)
Age	0.043	(148.1)	0.019	(58.9)
Age Squared[a]	-0.043	(-116.4)	-0.014	(-33.7)
Education	0.069	(469.4)	0.068	(398.8)
Part-time Job	0.093	(30.3)	0.104	(36.2)
Early-leaver	-0.001*	(-0.7)	-0.003*	(-1.6)
Size	0.059	(231.8)	0.055	(210.1)
Sales per Worker	0.101	(237.3)	0.083	(179.2)
OE (Out of Employment)	-0.066	(-19.2)	-0.041	(-11.5)
JC (Job Changes)	-0.014	(-2.9)	-0.014	(-2.9)
CIND (Changed Industry)	-0.067	(-18.9)	-0.127	(-33.1)
Tenure	0.007	(37.2)	0.013	(61.8)
Tenure Squared[a]	-0.003	(-5.0)	-0.024	(-36.6)
Constant	3.565	(556.7)	3.950	(593.3)
$\overline{R}^2$	0.60		0.62	
Number of Observations	773104		501494	

Notes: (i) a set of time, industry and regional dummies is included in the specification; (ii) [a] variables divided by 100;
(iii) all estimates are significant at 1%, except those with an *.

APPENDIX C - Selectivity Test

Table C.3.1: Sample Selection Model - Displaced Workers

	Reemployment Equation (REEMP=1)		Wage Equation Selectivity Correction		Wage Equation No Selectivity Correction	
	(1)		(2)		(3)	
Independent Variables	Coef.	t-ratio	Coef.	t-ratio	Coef.	t-ratio
Age	0.011	(3.8)	0.030	(21.6)	0.030	(21.5)
Age Squared[a]	-0.042	(-10.8)	-0.029	(-15.6)	-0.028	(-15.2)
Education	-0.005	(-2.7)	0.055	(77.6)	0.056	(78.6)
Tenure	-0.015	(-7.6)	0.035	(10.7)	0.035	(10.7)
Tenure Squared[a]	0.013*	(1.8)				
Female	-0.093	(-10.6)	-0.213	(-53.9)	-0.211	(-54.1)
Early-leaver	-0.481	(-43.2)				
Manufacturing	0.108	(11.9)	-0.001*	(-0.3)	-0.004*	(-0.8)
Log Hourly Earnings	-0.065	(-5.8)				
Log Unemployment Rate	-0.284	(-10.6)				
Part-time Job			0.085	(12.2)	0.086	(12.3)
Size			0.040	(37.9)	0.040	(38.0)
Sales per Worker			0.0006*	(1.1)	0.0007*	(1.5)
$Disp_1$			-0.043	(-7.3)	-0.046	(-7.9)
$Disp_2$			-0.056	(-7.5)	-0.061	(-8.3)
$Disp_3$			-0.048	(-4.8)	-0.056	(-5.7)
$Disp_4$			-0.065	(-4.6)	-0.076	(-5.6)
Constant	0.874	(10.3)	4.692	(165.1)	4.733	(181.7)
Time Dummies	no		yes		yes	
Regional Dummies	yes		yes		yes	
λ			0.050	(3.5)		
ρ			0.14			
Chi Squared; $\overline{R}^2$	5798.3		0.31		0.31	
Number of Observations	105507		33783		33783	

Notes: (i) REEMP: dummy variable equal to 1 if reemployed (0 otherwise); Female: 1 if female (0 if male); Manufacturing: 1 if employed in manufacturing (0 otherwise); Unemployment rate: aggregate (source - INE).
(ii) [a] variables divided by 100; (iii) all estimates are significant at 1%, except those with an *;

References

[1] Abraham, K. and H. Farber (1987), "Job Duration, Seniority, and Earnings", *American Economics Review*, Vol. 77, pp. 278-297.

[2] Abraham; K. and H. Farber (1988), "Returns to Seniority in Union and Nonunion Jobs", *Industrial and Labor Relations Review*, Vol. 42, pp. 3-19.

[3] Addison, J. and P. Portugal (1987), "The Effect of Advance Notification of Plant Closings on Unemployment", *Industrial and Labor Relations Review*, Vol. 41, pp. 3-16.

[4] Addison, J. and P. Portugal (1989), "Job Displacement, Relative Wage Changes, and Duration of Unemployment", *Journal of Labor Economics*, Vol. 7, pp. 281-302.

[5] Altonji, J. and R. Shakotko (1987), "Do Wages Rise with Job Seniority", *Review of Economic Studies*, Vol. 54, pp. 437-459.

[6] Arulampalam, W. (2001), "Is Unemployment Really Scarring? Effects of Unemployment Experience on Wages", *Economic Journal*, Vol. 111, pp. F585-F606.

[7] Bender, S., C. Dustmann, D. Margolis and C. Meghir (1999), "Worker Displacement in France and Germany", Institute for Fiscal Studies Working Paper N° 99/14.

[8] Blanchard, O. and P. Diamond (1994), "Ranking, Unemployment Duration and Wages", *Review of Economic Studies*, Vol. 61, pp. 417-434.

[9] Blanchflower, D., A. Oswald and M. Garrett (1990), "Insider Power in Wage Determination", *Economica*, Vol. 57, pp. 143-170.

[10] Bowlus, A. and L. Vilhuber (2002), "Displaced Workers, Early Leavers, and Re-employment Wages", *mimeo.*

[11] Burda, M. and A. Mertens (2001), "Estimating Wage Losses of Displaced Workers in Germany", *Labour Economics*, Vol. 8, pp.15-41.

[12] Carrington, W. (1993), "Wage Losses for Displaced Workers: Is It Really the Firm That Matters?", *Journal of Human Resources*, Vol. 28, pp. 493-496.

[13] Dustmann, C. and C. Meghir (1999), "Wages, Experience and Seniority", Institute for Fiscal Studies Working Paper N° 01/01.

[14] Fallick, B. (1996), "A Review of the Recent Empirical Literature on Displaced Workers", *Industrial and Labor Relations Review*, Vol. 50, pp. 5-16.

[15] Gregory, M. and R. Jukes (2001), "Unemployment and Subsequent Earnings: Estimating Scarring Among British Men 1984-94", *Economic Journal*, Vol. 111, pp. F607-F625.

[16] Hamermesh, D. (1989), "What Do We Know About Worker Displacement in the U.S.?", *Industrial Relations*, Vol. 28, pp. 51-60.

[17] Heckman, J. (1979), "Sample Selection Bias as a Specification Error", *Econometrica*, Vol. 47, pp. 153-161.

[18] Holmlund, B. and J. Zetterberg (1991), "Insider effects in Wage Determination - Evidence from Five Countries", *European Economic Review*, Vol. 35, pp. 1009-1034.

[19] Jacobson, L., R. LaLonde and D. Sullivan (1993), "Earnings Losses of Displaced Workers", *American Economic Review*, Vol. 83, pp. 685-709.

[20] Jones, S. and P. Kuhn (1995), "Mandatory Notice and Unemployment", *Journal of Labor Economics*, Vol. 13, pp. 599-622.

[21] Kletzer, L. (1989), "Returns to Seniority After Permanent Job Loss", *American Economic Review*, Vol. 79, pp. 536-543.

[22] Kletzer, L. (1991), "Earnings After Job Displacement: Job Tenure, Industry, and Occupation" in J. Addison ed., *Job Displacement: Consequences and Implications for Policy*, Detroit: Wayne State University Press, pp. 107-135.

[23] Kletzer, L. (1998), "Job Displacement", *Journal of Economic Perspectives*, Vol. 12, pp. 115-136.

[24] Kouch, K. (2001), "Earnings Losses and Unemployment of Displaced Workers in Germany", *Industrial and Labor Relations Review*, Vol. 54, pp. 559-572.

[25] Kruse, D. (1988), "International Trade and the Labor Market Experience of Displaced Workers", *Industrial and Labor Relations Review*, Vol. 37, pp. 557-569.

[26] Lefranc, A. (2002), "Labor Market Dynamics and Wage Losses of Displaced Workers in France and the United States", *mimeo.*

[27] Lehmann, H., K. Phillips and J. Wadsworth (2002), "The Incidence and Cost of Job Loss in a Transition Economy: Displaced Workers in Estonia, 1989-1999", IZA Discussion Paper N° 567.

[28] Lockwood, B. (1991), "Information Externalities in the Labour Market and the Duration of Unemployment", *Review of Economic Studies*, Vol. 58, pp. 733-753.

[29] Machado, J. and J. Mata (1998), "Earning Functions in Portugal 1982-1994: Evidence from Quantile Regressions", Banco de Portugal Working Paper WP 2-98.

[30] Margolis, D. (1999), "Part-Year Employment, Slow Reemployment, and Earnings Losses: The Case of Worker Displacement in France", in J. Haltiwanger *et al.* eds., *The Creation and Analysis of Employer-Employee Matched Data - Contributions to Economic Analysis*, North-Holland: Elsevier Science B.V., pp. 375-416.

[31] Nickell, S. and S. Wadhwani (1990), "Insider Forces and Wage Determination", *Economic Journal*, Vol. 100, pp. 496-509.

[32] Phelps, E. (1972), *Inflation Policy and Unemployment Theory*, London: Macmillan.

[33] Pissarides, C. (1992), "Loss of Skill during Unemployment and the Persistence of Employment Shocks", *Quarterly Journal of Economics*, Vol. 107, pp. 1371-1391.

[34] Podgursky, M. and P. Swaim (1987), "Job Displacement and Earnings Loss: Evidence from the Displaced Worker Survey", *Industrial and Labor Relations Review*, Vol. 41, pp. 17-29.

[35] Rosolia, A. (2002), "The Consequences of Job Displacement in Italy", *mimeo.*

[36] Ruhm, C. (1991), "Are Workers Permanently Scarred by Job Displacements?", *American Economic Review*, Vol. 81, pp. 319-324.

[37] Ruhm, C. (1992), "Advance Notice and Postdisplacement Joblessnes", *Journal of Labor Economics*, Vol. 10, pp. 1-32.

[38] Ruhm, C. (1994), "Advance Notice, Job Search, and Postdisplacement Earnings", *Journal of Labor Economics*, Vol. 12, pp. 1-28.

[39] Stevens, A. (1997), "Persistent Effects of Job Displacement: The Importance of Multiple Job Losses", *Journal of Labor Economics*, Vol. 15, pp. 165-188.

[40] Swaim, P. and M. Podgursky (1990), "Advance Notice and Job Search: The Value of an Early Start", *Journal of Human Resources*, Vol. 25, pp. 147-178.

[41] Topel, R. (1991), "Specific Capital Mobility, and Wages: Wages Rise With Job Seniority", *Journal of Political Economy*, Vol. 99, pp. 145-175.

[42] Wooldridge (2002), "Inverse Probability Weighted M-Estimators for Sample Selection, Attriction, and Stratification", *Portuguese Economic Journal*, Vol. 1, pp. 117-140.

Chapter 4
Wages and the Risk of Displacement

1 Introduction

The extent of job destruction and, in particular, firm closing and job loss due to sector reallocation, has been a matter of great concern in recent years, with empirical research on gross job flows experiencing a tremendous growth in the past decade. The studies on the decomposition of net employment flows emphasize the importance of job creation and job destruction through the entry and exit of firms. According to Davis *et al.* (1996), about one-fourth of annual job destruction in the U.S. takes places at plants that shut down, while startups account for one-sixth of annual job creation. In Portugal, annual job flows produced by both plant births and plant deaths account for almost half of total gross employment flows (see Blanchard and Portugal, 2001).

However, the literature on flows of jobs is mostly employment accounting, whereas wages/prices and fluctuations in labor demand are never considered. As pointed out by Hamermesh (1993): "...data on gross flows of jobs tell us nothing directly about the magnitude of the wage or output elasticities of employment changes through the births and deaths of establishments, or growth or contraction in existing establishments."

Recently, a considerable number of studies have been examining how the wages of displaced workers vary (over a long-term period) compared to workers who are not displaced [see, among others, Jacobson *et al.* (1993), Stevens (1997) and Margolis (1999)]. Nonetheless, few studies have yet analyzed how that wage variation affects the probability of displacement. In fact, the theoretical and empirical research on the role of wages on plant closings is remarkably sparse. Most of the empirical literature on plant closings has been concentrated on the effect of unions in the probability that a firm (plant) shuts down.[1]

Hamermesh (1988) was the first who explicitly addressed this issue, developing a model in which workers and firms contract over wages and employment probabilities. Since then, and to the best of my knowledge, empirical research in this area has not seen great improvement. The exceptions are the studies of Dunne and Roberts (1990) and Blanchflower (1991).

Hamermesh (1988, 1991 and 1993) modeled the relationship between wage changes and the probability of job displacement due to plant closing, in order to determine the necessary wage concessions to keep plants from closing. The

[1] See Addison *et al.* (2002, pp. 23-24) for a summary of the international evidence of union effects on plant closings for Britain and the Unites States. Their own study is about the effects of worker representation on plant closings in Germany.

model estimates also allowed him to compute the elasticity of labor demand through plant closings.

The model is set within a theoretical contract framework in which workers contract with their employers for a package that includes a probability that the job will exist and a wage premium above the entry-level wage (reservation wage). In fact, when workers sort themselves among firms, one of the risks they consider is that exogenous product-market shocks may cause the firm to close down. From the perspective of the contractual relationship between employers and employees the role of joint investments in specific training may be viewed as a buffer that can cushion against negative shocks and, thus, partially insulate the firm from unfavorable market conditions. That is, since an internal labor market may operate with employers and employees sharing the rents originated by firm-specific human capital, the adjustment to negative shocks may be partially absorbed through wage concessions. The possibility of wage concessions is, of course, precluded if workers are paid legal minimum wages.

Hamermesh's model has two main predictions. The first points to a negative relationship between the excess of wages over the reservation wage and the probability of closing, suggesting that shocks that increase the probability of displacement reduce the magnitude of the wage increase. The second points to the existence of a positive relationship between the reservation wage and the probability of closure due to the existence of compensating differentials for the *ex ante* risk of displacement.

In order to test these predictions, Hamermesh (1998, 1991) used a longitudinal household sample from the Panel Study of Income Dynamics (PSID) that includes 114 workers who left their previous job between 1977-81 because the plant shut down, and 2,433 household heads who were not displaced during this period. His results revealed that shocks that increase the probability of displacement also significantly reduce the wage increase. In other words, average wages grow less rapidly in plants that will soon close, suggesting that firms' adjustment to price shocks are partially absorbed into wages. Concerning the second prediction of the model, no robust evidence was found in favor of the existence of compensating differentials for the *ex ante* risk of displacement. In fact, and contrary to what was suggested, a negative coefficient was found for the reservation wage, a finding that Hamermesh interpreted as reflecting the dominance of the income effect of higher earnings on the demand for security.

Dunne and Roberts (1990) used a simultaneous equations approach to examine how the probability of plant closing affects the wage paid to employees and how wages affect the probability of closure. For this purpose they develop a two-equation empirical model of plant failure and wage determination in order to estimate both the compensating differential employees required for the risk of plant closing and the effect of wage changes on the probability of plant failure.

Using longitudinal data on over 6,500 manufacturing plants from the U.S. Annual Survey of Manufactures for the years 1974-78, they found evidence in favor of the existence of compensating differentials for the risk of displacement due to plant closing. In fact, workers in a plant that has the average probability of failure in the sample, earn 7.3% higher wages than workers with a zero prob-

ability of failure. This result, however, remains robust only for plants owned by multiplant firms. Finally, the results revealed that wage increases have little effect on the probability of plant failure. Indeed, a ten percent increase in wages, holding plant revenue fixed, increases the probability of failure by only 0.15 percentage points.

Using microdata on 5,300 individuals from the British Social Attitude Survey between 1983 and 1989, Blanchflower (1991) analyzed how the risk of unemployment may affect wages. He added to a classic Mincerian cross-section wage equation a range of variables related to the extent of excess supply in the labor market. One of these variables is a proxy for the risk of displacement due to plant closure. Using a Nash bargaining framework he demonstrated that wages should be a declining function of the probability of firm closure.[2] The empirical estimates for the full sample revealed that workers who reported that they expected their plant to close earn, on average, 8% less, *ceteris paribus*. This result seems to suggest that, if anything, fear of job loss will generate bargaining concessions instead of compensating differentials.

Some case studies also showed that unions may agree to moderate wage demands if jobs are threatened, and if such moderation is likely to generate a clear improvement in employment security [see, for example, Cappeli's (1985) case study in the meatpacking and tire industries]. As mentioned by Cappeli, the possibility of shutdowns at the firm level may be a particular threat to employment security, because seniority systems that usually protect most workers from layoffs generally do not offer protection against firm closings. In the face of a firm shutdown insider workers are all equal.

In sum, the empirical research on the effect of wages on firm closings (and *vice versa*) is still in the beginning and looking for a solid stylized fact concerning the relationship between wages and the risk of unemployment. Do workers require a compensating differential for the risk of displacement? Or, on the contrary, are they able to accept wage moderation in order to avoid the firm's shutdown? Are wages an important determinant of firm exit?

Using a nationally representative data set that links employers and employees, this study will examine how wages adjust to a negative demand shock that raises the risk of displacement through firm closing and to what extent a wage change affects the exit likelihood. The role of a mandatory minimum wage on the firm's exit decision will also be analysed.

This work attempts to give an additional contribution to the empirical literature on wages and the risk of displacement on three distinct grounds. The first is related to the use of an appropriate and representative data set to analyze the relationship between wages and the risk of displacement. In fact, the Portuguese data from *Quadros de Pessoal* (*QP*) can be described as a longitudinally matched worker-firm sample with a rich set of information on workers' characteristics, their wages, and their work environment. This will enable us to address a number of questions that cannot be adequately answered in the absence of firm or worker data.

[2] For a demonstration of this prediction see Blanchflower (1991, Appendix A, pp. 492-494).

The second is related to the use of a simultaneous-equations approach in order to account for the possible endogeneity of wages and the probability of displacement. As it seems clear that an increase in the firms' failure rate may affect wage changes because it raises the risk of unemployment, it also seems clear that a wage change may affect the exit likelihood because it reduces, all else being equal, firm's profitability.

Third and finally, this study makes an important contribution by examining the effect of a mandatory minimum wage on the failure rate. Despite the great effort dedicated to research on the effect of minimum wages on unemployment, namely, youth unemployment, we are not aware of any study that explicitly looks at the relationship between minimum wages and firms' exits. Are firms with a higher proportion of mimimum wage earners more vulnerable to product shocks due to their inability to adjust wages downward?

The plan of the paper is as follows. Section 2 presents the simultaneous-equations model of firm closing and wages. In this Section we also develop the basic hypothesis regarding which factors should matter for exit and discuss the wage determinants. In Section 3 the data set is described. Section 4 reports the empirical results and Section 5 concludes.

2 The Empirical Model of Firm Closing and Wages

2.1 Purpose

The empirical model of firm closing and wages presented in this Section attempts to tackle three different questions. The first is to examine how wages adjust to a negative demand shock that raises the probability of displacement through firm closing. The second is to determine how wages themselves affect the exit likelihood. The third is to analyze the effect of a mandatory minimum wage on the failure rate.

With respect to the first issue, the objective is to examine if a higher risk of displacement may lead workers to accept wage concessions or if, on the contrary, that threat leads workers to demand a compensating differential. In fact, in the face of a risk of layoff two forces may be at work. First, workers may demand higher wages in order to compensate them for the higher risk of displacement. This is the prediction of the competitive model with its roots in Smith (1974). Fear of unemployment has to be compensated, like any other disutility, with a wage premium. Second, workers may accept wage moderation (concessions) in order to avoid the firm closing and subsequent displacement. This hypothesis is consistent with the idea that pay is fixed in a bilateral bargain where the fear of unemployment acts to weaken workers' bargaining position. If the first force prevails a positive correlation between wages and the probability of plant closing should be expected. Conversely, if it is the second force that prevails, a negative correlation should be expected.

Concerning the second issue, the effect of wages on firm's profitability, and consequently on firm's survival, is analyzed. The effect of wage levels in the probability of firm closing may be ambiguous. One would expect that all else being equal, firms with lower wages would have higher expected profits, and thus be more likely to survive. However, if higher wages are compensated by increased worker quality or raises in worker productivity, as suggested in the efficiency wage literature [Katz (1986)], there might be no correlation.

In the third issue, particular attention will be devoted to the role of minimum wages on firms' closure. In fact, the adjustment to negative demand shocks may be partially absorbed through wage concessions. Nonetheless, the possibility of wage concessions is precluded if workers are paid legal minimum wages. Thus, we examine if a mandatory minimum wage imposes severe restrictions in wages adjustment to unfavorable demand schocks that may accelerate the firm's exit decision.

At this point the reasons that led to choosing a model of firm closing and not plant closing shoul be mentioned. The option to use information at the firm level instead of at the plant level is justified by two main reasons. First, the important management bargaining decisions in a multiplant firm are made at the corporate level, not at the plant, and reflect the priorities of the firm as a whole. In particular, wage policies are mainly relevant at the firm level. Second, it seems that when it is the firm that is at risk of closing, the unemployment threat is stronger than when it is an establishment of a multiplant firm. In large multiplant firms plant shutdowns may be used in addition to layoffs as a means of reducing capacity in the face of unfavorable shocks in the product demand. A plant shutdown may even occur with no layoffs, as workers from closing plants are reemployed in other plants of the same firm. Indeed, in some situations the shutdown of a plant may be a less costly bargaining strategy and, thus, would be preferable to a wage concession strategy. The empirical evidence for Portugal suggests that due to higher adjustment costs (mainly the costs of firing workers) and in the face of unforseen temporary shocks, it is preferable to employers, under certain circumstances, to close down instead of adjusting their level of employment by laying off workers (Blanchard and Portugal, 2001).

2.2 The Empirical Model

As mentioned above, the three main objectives of the empirical model of firm closing and wages are:

(i) to examine how wages in period *t-1* are affected by a higher risk of displacement due to a threat of firm closure in period t;

(ii) to examine the relationship between the wage paid to an individual worker in the last year on the job (period *t-1)* and the probability that the firm closes in the next year (period t);

(iii) to analyze whether a higher incidence of minimum wage earners in period *t-1* affects the firm's failure rate in period t.

The basic model consists of two equations. The first describes the probabilistic event of a displacement due to plant closure, and is specified from the following equation:

$$\pi^*_{ijt} = \alpha_1 X_{jt-1} + \alpha_2 Z_{ijt-1} + \alpha_3 W^*_{ijt-1} + \alpha_4 \Phi(W_{Mit-1}) + \upsilon_{1ijt-1} \quad (4.1),$$

$$\text{where } Y_{ijt} = 1 \text{ if } \pi^*_{ijt} < 0 \text{ and } Y_{ijt} = 0 \text{ if } \pi^*_{ijt} \geq 0 \quad (4.2).$$

π^*_j is a latent variable reflecting the future profitability of firm j. X_j is a vector of firm characteristics, Z_{ij} a vector of worker characteristics and W_{ij} the natural logarithm of the wage paid to worker i by firm j. $\Phi(W_{Mi})$ denotes the probability of a given worker receiving the minimum wage. α's are the parameters to be estimated and υ_{1ij} is a normally distributed random variable with zero mean and unit variance.

The second equation of the model is a conventional human capital wage equation with controls for local labor market conditions and some firm characteristics. Generically, the wage equation is defined as:

$$W^*_{ijt-1} = \beta_1 U_{jt-1} + \beta_2 V_{ijt-1} + \beta_3 \pi^*_{ijt} + \upsilon_{2ijt-1} \quad (4.3),$$

where $W_{ijt-1} = Max(W_{Mit-1}, W^*_{ijt-1})$ and W_{Mit-1} is the mandatory minimum wage in period $t-1$.

The wage paid to worker i in firm j is a function of a set of workers' characteristics included in vector V_{ij}, firm and local labor market characteristics defined in vector U_j and π^*_{ijt}.

In the data it is not possible to observe π^*_{ijt}. All we can say is whether π^*_{ijt} is or is not below a given threshold (the minimum level of profits that guarantees the firm's continued existence). In the latter case, the firm will continue its operations, otherwise it will close down. Thus, the probability of displacement through firm closing is defined as $P_{ijt} \equiv \Pr(\pi^*_{ijt} < 0)$. β's are the unknown parameters to be estimated and υ_{2ij} is a normally distributed random variable (zero mean and constant variance).

The dependent variable in the wage equation is defined as the natural logarithm of the real monthly base wage paid to an individual worker in the year that precedes the displacement. The monthly base wage was deflated by the Consumer Price Index (CPI; base=1991).

The wage equation includes a set of controls for personal characteristics such as: gender, education, age (and its square), tenure (and its square) and qualification level. The variable gender takes the value one for females and zero for males. The variable education is measured as the number of years of shooling completed and tenure is defined as the number of years with the current employer. A set of dummies are used for the levels of qualification. Seven categories are considered: manager and highly professional, professional, supervisors, highly skilled and skilled, semi-skilled and unskilled, non-defined (a residual category) and apprentices (the reference category).

In order to assure that the effect of a higher risk of unemployment on wages is due to firm shutdown and not to differences in the risk of layoff in the local labor markets, the local unemployment rate is used to control for those differences. The local unemployment rate is defined at the disaggregated level of NUTs III.[3]

The firm characteristics included in the wage equation are the firm size, sales per worker and a set of industry (one-digit)[4] and regional (NUTs II)[5] dummies. Since the data include firm closures that occurred in 1994, 1995 and 1996, three time dummies were also added to the model in order to control for macroeconomic conditions.

The dependent variable in the failure equation is a binary variable that takes the value of one if the worker was displaced in year t due to firm closure, zero otherwise. A set of firm variables that may affect the firm's decision to close are identified below.

In order to control the exogenous demand shocks that may affect the probability of firm closing, the average growth rate of real sales in the last three years is used as a proxy for firm-specific skocks.[6] Controlling for firm-specific (idiosyncratic) shocks enables one to examine if, in the face of an identical exogenous shock, and all else being equal, firms with lower wages are less likely to close down or not.

While firm or sectoral demand shifts certainly affect the rate of firm closing, other forces are also at work. Jovanovic (1982) showed that patterns of plant and firm failure are consistent with a process of within-industry selection in which relatively high-cost producers fail. This selection process leads to substantial variation in the probability of failure across firms (plants) within an industry. In fact, while plant deaths are part of the normal process of the entry and exit of firms, there is substantial heterogeneity in the probability of shutdown across plants, firms and industries (Dunne *et al.*, 1989). This reasoning suggests that the risk of displacement faced by workers as a result of firm closing varies not only with the demand for the firms' output but also with the efficiency of their employer relative to competing firms in the same industry. Being so, a set of firm's characteristics that are related to its performance in the output market may affect the firm's own probability of survival. The factors to be included in the empirical model as exit determinants will now be identified.

[3] It should be noted that for the period of analysis of this study (1993-95), unemployment rates are only defined at the regional level at NUTs II. In order to have a proxy for unemployment at a more disaggregated level of NUTs III (28 geographical areas for mainland), the ratio between annual job applications registered in each employment center and total employment (defined at NUTs III using data from *QP*) will be used.

The information on job applications registered in each employment center was obtained from *Monthly Statistics - Institute for Employment and Vocational Training (IEFP).*

[4] At one-digit level there are nine sectors according to the Portuguese Classification of Economic Activities (CAE). In this case we employ just 7 sectoral dummies, since the sector of Electricity, Gas and Water, that accounts for a few number of non displaced workers, was excluded because no plant closings were observed in this sector. Including this sector will not enable to obtain reliable estimates for the coefficient of this sector in the probit equation.

[5] At NUTs II mainland Portugal is split into 5 geographical areas.

[6] Sales in year t correspond to annual sales of the previous year. Sales were also deflated using the CPI (base=1991).

With the exception of past sales growth, all variables are measured in the year that precedes the exit event (period *t-1*).

It has been largely shown in the empirical literature on firm survival that firm size and age are negatively associated with failure rates [see, for example, Kumar (1985), Evans (1987), Hall (1987), Dunne *et al.* (1989), Audretsch and Mahmood (1994), Mata and Portugal (1994) and Mata *et al.* (1995)]. These results are consistent with Jovanovic's (1982) model of industry evolution, according to which firms start with no knowledge about their efficiency. As time goes by and firms observe their performance in the output market, they gradually learn about their efficiency. This information is then incorporated into their current size. Efficient firms grow and survive, while inefficient ones contract and fail. Thus, large and old firms are successful firms, and, for this reason, they should have higher survival probabilities.

For this reason, measures of the size and age of the firm are included in the failure equation. The size of the firm is defined as the natural logarithm of its total employment. Since the information about the date of a firm's creation is only available after 1993, and in order to use the same criteria to measure age in the 1993-95 period, we used as a proxy for firm age the tenure (in years) of the worker with the longest tenure within the plant. A linear spline function is used to define the variable age of the firm.

The firm's market share is used as a measure of product market competition. Monopoly power generates monopoly rents and consequently higher profits. If employers are able to appropriate part of these rents, it should be expected that firms with increased market share are less likely to fail. The market share is obtained by the ratio between a firm's sales and total (5 digit) sector's sales.

Firm ownership characteristics may affect the exit likelihood. Two indicators of ownership type will be used, namely the number of establishments with which each firm operates and the proportion of foreign capital. For the former a dummy variable that takes the value one if the firm is a multiplant firm (0 otherwise) will be included in the model. Single-plant firms are far less likely to shutdown a plant than a multi-plant firm, since single-plant operations can be viewed as having greater closing costs as they involve the exit of the own firm. Therefore, plants in multi-unit firms have higher probability of closing. The empirical evidence has been showing that multi-plant firms use the shutdown margin more often than their single-plant counterparts [see, for instance, Mata and Portugal (1994), Machin (1995) and the recent studies of Addison *et al.* (2002) and Bernard and Jensen (2002)]. On the contrary and analogously, multi-plant firms are far less likely to close than single-plant firms.

The proportion of foreign capital may itself be an indicator of unobserved quality of the firm and may affect the probability of closure. Doms and Jensen (1998) found that multinational plants have superior observable characteristics. However, it is also well known that multinationals have a higher propensity to relocate production within firms, which may lead to an increased probability of closure [see, for example, Harris and Hassaszadeh (2002)].

If analyzing the effects of individual wage levels on the probability of firm closing, it is necessary to have a measure of revenue per employee in order to be

able to compare firms. Firms that have higher variable costs, holding revenue fixed, are less likely to cover their fixed costs in the long run and thus more likely to close down. Real sales per employee (in logs) is used as a measure of firm's revenue per employee.

Finally, among the firm's characteristics, a set of industry (one-digit) and regional dummies (NUTs II) are also included in the failure equation. In order to account for macroeconomic conditions time dummies are also added to the model.

Previous studies such as Cooper *et al.* (1994) and Mata and Portugal (2002) found human capital to be a good predictor of survival. We also consider the possibility that human capital affects the firm's performance and hence its failure rate, and include a set of variables that characterize workers' skills and that identify the composition of the firm's workforce with respect to schooling, tenure (and its square), age and gender. The variable age is defined as a dummy variable that equals one if the individual belongs to each age group out of the four considered: less than 25, 25-34, 35-54 and more than 54 (the omitted category).

3 The Data

The data set used in this study was obtained from *Quadros de Pessoal (QP)* and includes all workers that lost their jobs in 1994, 1995 or 1996 due to firm closure and were present in the *QP* registers in the year that preceded the displacement. A control group constituted by a random sample of workers who were employed in the year prior to the displacement in firms that did not close in the next year is also included.

The survey has three characteristics that make it particularly suitable for the analysis of the relationship between wages and the risk of firm closing. First, it covers all firms employing paid labor in Portugal.[7] Second, it has a longitudinal dimension which allows us to follow firms and individuals over time. Third, it contains information on both firms and its workers.

To be sure that exits are accurately identified, we required that a firm be absent from the *QP* two or more consecutive years in order to be classified as a closure. Additionally, and to avoid the inclusion of false exits, workers that appeared in the database in the period after displacement with a year of admission in the new job less than the year of displacement minus one were dropped.[8]

In order to be allowed to construct the variables that account for firm's recent evolution, we impose that workers be present in the *QP* registers in each

[7] Thus, this source does not cover operated family businesses without wage-earning employees and self-employment. Public administration is also excluded.

[8] If, for example, a worker's displacement year is 1994 and he (she) appears in the database in the post-displacement period with a year of admission in the new job of 1992 or less, he (she) is excluded from the sample.

of the three years that preceded the firm shutdown and employed with the same employer over those years.[9] This requirement means that an individual must have at least two years of tenure in the year prior to displacement. This selection rule, although primarily dictated by data availability considerations, results in an analysis sample of stably employed individuals. On average, the sampled individuals have 26 years of labor force experience and over 11 years of employer tenure.

We also limited the sample to full-time workers aged between 18 and 64 in the year prior to displacement. Since the minimum wage is defined as a monthly wage, the full-time job requirement allows to identify minimum wage earners more accurately. In this context, wages are measured as monthly wages.

We have also excluded those individuals for which information was incomplete for the year before displacement, namely those with zero wage.[10] Finally, and in order to avoid the outliers in wages and sales, we drop the 0.1% top and the 0.1% bottom observations for those variables.

After these exclusions we obtained a sample of 35,922 full-time workers that were displaced between 1994-96 due to firm closing, aged between 18-64 and with at least two years of tenure in the year prior to displacement.

The control group includes three sub-samples and was constructed in the following way. For each year prior to the displacement year we obtained a random sample of around 300,000 workers that were employed in firms that did not close.[11] For each of these three groups we excluded those individuals that were not present in the *QP* files in each of the three years before displacement and those who were not employed in the same firm over those years. The sample was also limited to full-time workers aged between 18-64 in the year prior to displacement. After excluding those observations with missing values on the explanatory variables and the outliers in wages and sales, we obtained a control group of 230,102 non-displaced workers.

Table 4.1 presents the descriptive statistics of the sample for the two groups of workers: displaced and non-displaced. As can be seen in Table 4.1, on average, displaced workers are slightly younger, less qualified and with fewer years of tenure and education. They also earn, on average, less than non-displaced workers. The pool of displaced workers includes more females and minimum wage earners.

According to firms' characteristics, the proportion of displaced workers that comes from small, young and single-plant firms is higher when compared to the sample of non-displaced workers. For the former the real average growth rate of firms' sales in the last three years is negative (-7.6%), while for the latter that same rate is positive (1.4%). Displaced workers are also employed in firms with a reduced market power (measured by market share).

[9] Hence, for workers displaced in 1994 the data should be available for the 1991-93 period, for workers displaced in 1995 for the period of 1992-94 and for workers displaced in 1996 for the 1993-95 period.

[10] For the three years before displacement in the case of the variable sales.

[11] The sample was drawn according to a normal random number generator.

Table 4.1: Sample Characteristics (Means and Standard Deviations)

	Displaced		Non-displaced	
	Mean	St. Dev.	Mean	St. Dev.
Workers' Characteristics				
Age (in years)	37.3	0.113	38.5	0.107
Tenure (in years)	9.8	0.082	11.9	0.085
Education (in years)	5.8	2.830	6.4	3.231
Proportion of Female	0.442		0.368	
Proportion of Minimum Wage Earners	0.143		0.056	
Qualification Levels (proportion of workers)				
Manager and Highly Professional	0.018		0.029	
Professional	0.017		0.031	
Supervisors	0.051		0.057	
Highly Skilled and Skilled	0.527		0.543	
Semi-skilled and Unskilled	0.279		0.262	
Apprentices	0.057		0.034	
Non-defined	0.051		0.044	
Firms' Characteristics				
Size (total employment)	108.4	426.2	1135.1	2701.2
Past Sales Growth	-0.076	0.391	0.014	0.354
Market Share	0.013	0.060	0.118	0.250
Proportion of Foreign Capital	0.031	0.161	0.094	0.269
Proportion of Multi-Plant Firms	0.165		0.413	
Firm Age (proportion of workers)				
2-5 years	0.134		0.046	
6-10 years	0.181		0.086	
$>$ 10 years	0.685		0.869	
Real Sales per Worker (in logs)	8.419	1.111	8.825	1.172
Real Monthly Wage (in logs)	10.973	0.432	11.175	0.490
Number of Observations	35922		230102	

Notes: (i) all variables, except past sales growth, are measured in the year prior displacement; (ii) sales per worker and the monthly wage are in 1991 PTE (escudo); 1 EUR≡200.482 PTE.

4 Estimation and Empirical Results

4.1 Estimation Method

In order to estimate the simultaneous-equations model of firm closing and wages presented in Section 2, it will be necessary to use an appropriate method of estimation. It is well known that the ordinary least squares (OLS) method is generally biased and inconsistent when applied to a structural equation in a simultaneous-equations system.

Beyond this difficulty, the empirical model of firm closing and wages has a particularity, since one of the endogenous variables is a binary variable while the other is a censored variable. In fact, while the failure equation is specified as a probit model, the wage equation is a tobit model with lower censoring at the minimum wage. Thus, we are in the presence of a simultaneous-equations model with mixed dichotomous and censored variables.

The conventional method for estimating simultaneous-equations models is the method of instrumental variables. As suggested by Maddala (1983, p. 246), a two-stage procedure will allow us to estimate a two-equation model in which one of the variables is censored while the other is only observed as a dichotomous variable. The two-stage estimation method involves the following steps. The first step is to write the reduced-forms equations for the endogenous variables. Next, estimate the reduced-forms equations and keep the predicted values. Finally, estimate the structural equations replacing the endogenous variables by the predicted values obtained from the reduced-forms regressions.

The empirical model defined by equations (4.1) to (4.3) specifies both the probability of closure and wages as endogenous. Hence, the reduced form of the equation system in the latent variables is:

$$\pi^*_{ijt} = \Pi_1 K_{ijt-1} + \varepsilon_{1ijt-1}, \qquad Y_{ijt} = 1(\pi^*_{jt} > 0) \tag{4.4},$$

$$W^*_{ijt-1} = \Pi_2 K_{ijt-1} + \varepsilon_{2ijt-1}, \qquad W_{ijt-1} = Max(W_{Mit-1}, W^*_{ijt-1}) \tag{4.5}.$$

where K includes all the exogenous variables in X, Z, U and V.

The reduced-form parameters can be estimated by applying maximum likelihood (ML) to the probit and tobit models in equations (4.4) and (4.5), respectively.

Estimating the reduced-form for π^*_{ijt} by the probit method will allow us to obtain the predicted probability of displacement through firm closing, $\widehat{P}$. Estimating the reduced-form for W^*_{ijt-1} by the tobit method will enable us to obtain the predicted value of the monthly wage ($\widehat{W^*}$) and the estimated probability that a given observation is a limit observation, $\Phi(\widehat{W_M})$. In other words, $\Phi(\widehat{W_M})$ measures the estimated probability of a given worker receiving the minimum wage.

This procedure is unconventional, but provides a simple and elegant solution to the specification of the two sources of endogeneity from wages to failure rates. On the one side, the impact of the level of wages on the chances of firm closure.

And, on the other side, the influence of minimum wage restrictions on the ability to accomodate negative shocks.

The structural wage equation is estimated in the second-stage tobit after replacing the probability of displacement through firm closing (P) by its predicted value ($\widehat{P}$). The structural failure equation is estimated in the second-stage probit after replacing the monthly wage (W^*) by its predicted value ($\widehat{W^*}$) and after including the estimated probability of being a minimum wage earner, $\Phi(\widehat{W_M})$. This last procedure will enable us to examine the effect of a mandatory minimum wage on the failure rate.

4.2 Empirical Results

The parameter estimates of the simultaneous-equations model of firm closing and wages are presented in Tables 4.2 (structural failure equation) and 4.3 (structural wage equation).[12] The estimation strategy consists of having, as far as possible, a complete set of controls to examine whether a robust association between wages and the probability of firm closing (and *vice-versa*) can be identified. The variables that characterize employers' performance such as past sales growth, firm age, market share, multi-plant firm and proportion of foreign capital contribute to the identification of the wage equation.[13] In the structural probit model of firm closing, the regional unemployment rate and the qualification levels are omitted.

Columns 1 and 2 of Table 4.2 report results (coefficients estimates and marginal effects, respectively) for a specification in which the probability of firm closing depends on an extensive set of firm characteristics, the skill composition of the workforce, monthly wages (predicted) and the estimated probability of being a minimum wage earner.

According to Table 4.2, past sales growth, firm size and age, market share, multi-plant firm, proportion of foreign capital and sales per worker are significantly correlated with the probability of firm closing. In particular, the results reveal that firms experiencing a decline in sales growth are clearly more likely to close. This seems to imply that sales contraction can be used as a strong predictor of firm failure. Indeed, the fact that a firm has grown in the past signals that it has been performing well. Moreover, the estimates reported in Table 4.2 show that small firms are clearly more likely to close than large firms. This result is conventional enough and, in particular, is in line with the one obtained for Portugal in the study of Mata *et al.* (1995) using a sample of newly born manufacturing plants.

The estimates of the coefficients on firm's age using splines indicate a negative and significant effect of age on the probability of displacement. However,

[12] The econometric results were obtained using LIMDEP version 8.0. In both equations the t-ratios correspond to the corrected covariance matrix for the two-step estimator using the methodology developed by Murphy and Topel (1985).

[13] The exceptions are firm size and sales per worker, since these variables constitute an important determinant of individual wages as well.

after a decade the negative effect of age starts to vanish, becoming positive for very old firms (more than 53 years).

The variable market share has a strong negative effect on the probability of closing, suggesting that monopoly power generates rents that may function as a buffer that cushions against negative shocks.

As expected, workers that are part of a multi-plant firm are less likely to be displaced due to firm closing than workers that are part of a single-plant firm. The same is true for workers that are part of firms with a large proportion of foreign-owned capital.

Sales per worker, a proxy for productivity, have a negative impact on the probability of firm closing. Thus, low productivity firms, all else being equal, are more likely to close down.

Concerning the variables that identify the composition of the firm's workforce, the estimates show that firms with a higher proportion of female, older, less-educated and less-tenured workers are more likely to close. In particular, we should mention the negative effect of tenure on the probability of closure. As long as higher tenure reflects higher investments in specific training and/or better matches (that is, job matches that enhance productivity), joint investments in specific training may also be viewed as a buffer than can cushion against negative shocks and, thus, partially insulate the firm from unfavorable market conditions. Hence, a negative relationship should be expected between worker's tenure and the probability of firm closing. In this context, we interpret the positive coefficient estimate associated with the gender variable as an indication that females engage in on-the-job training investments less intensively than do males.

High-wage paying firms face higher hazard rates than low-paying firms, *ceteris paribus*. After controlling for an extensive set of employers' characteristics and for the skill composition of the workforce, the results reveal that firms that pay higher wages, holding revenue per employee fixed, are less likely to survive. In fact, the marginal effect of a 1% increase on monthly wages in the probability of displacement is 0.00029 (see column 2 of Table 4.2). Since the average job displacement rate through firm closing in the population is around 6.3%, a 1% wage increase is associated with a 0.46% increase in the probability of job displacement through firm closing.

Finally, the two-step probit results report a positive and significant effect of the probability of receiving the minimum wage on the failure rate, suggesting that firms with a higher incidence of minimum wage workers face higher exit rates than those with a smaller incidence. A one point increase in the proportion of minimum wage earners increases the probability of displacement by 0.014 percentage points. Since, on average, the proportion of minimum wage earners in the population is around 13% and the average job displacement rate is 6.3%, a 1% increase in the proportion of minimum wage earners increases the probability of displacement through firm closing by 0.028%.

In fact, the possibility of wage concessions is precluded if workers are paid legal minimum wages. Thus, firms with a higher proportion of minimum wage earners may have lower chances of survival due to their inability to adjust wages

downward in the face of a negative demand shock.

Table 4.3 reports the two-step tobit results of the wage equation. The basic specification includes a set of human capital variables, firm size and sales per worker as two important determinants of individual wages, the local unemployment rate and the instrumented probability of displacement due to firm closing. All the exogenous variables are statistically significant at the 1% level of significance and have the expected signs.

The effect of the probability of closing on monthly wages is negative and also statistically significant. This implies that a worker employed in a firm that will close earns less in the year prior to displacement than a similar worker employed in a non-closing firm. Converting the coefficient of 0.459 to an elasticity results in a value of 0.029 evaluated at the mean failure rate in the sample.[14] In other words, workers in a firm that has the average probability of failure in the population (6.3%), earn (one year prior to closing) around 3% less than workers in a firm with zero probability of failure. This empirical result indicates the existence of wage concessions for workers employed in firms with a higher probability of failing.

This empirical evidence seems to contradict the theoretical prediction that workers employed in firms with a higher probability of displacement require a compensating differential for the risk of layoff. The compensating differential for the *ex ante* risk of displacement may exist. The evidence of a positive relationship between wage levels and failure rates reported in the probit equation is consistent with the existence of higher wages due to compensating differentials, that may accelerate in the short-run the process of closure.[15] However, the results obtained from the wage equation seem to suggest that in the face of an *ex post* risk of displacement the compensating differential for the risk of layoff is offset by the need to moderate wages in order to avoid the firm's shutdown.

[14] If $\widehat{W}_1$ is the predicted wage in a firm with an average failure probability of 0.135 and $\widehat{W}_0$ is the predicted wage in a firm that has no probability of failing, then relative wages are calculated as $\widehat{W}_1/\widehat{W}_0 = \exp((0.459) * (0.063)) = 1.029$.

[15] Since in the basic specification of Table 4.2 a control for tenure in the job was included, we can interpret the individual monthly wage as the worker's reservation wage.

Table 4.2: Failure Equation - Two-step Probit Results
Full-time Workers (N=266024)
Dependent variable: Displaced=1

Variables	Coefficient	Marginal Effect
Past Sales Growth	-0.323	-0.049
	(-30.5)	
Firm Size	-0.251	-0.038
	(-75.5)	
Firm Age		
Age	-0.056	-0.008
	(-5.0)	
AgeS5	-0.007*	-0.001
	(-0.5)	
AgeS10	0.069	0.010
	(23.0)	
Market Share	-1.383	-0.208
	(-28.1)	
Multi-plant Firm	-0.069	-0.010
	(-7.2)	
Proportion of Foreign Capital	-0.245	-0.037
	(-13.3)	
Sales per Worker	-0.065	-0.010
	(-14.8)	
Female	0.135	0.021
	(14.1)	
Education	-0.012	-0.002
	(-5.5)	
Worker Age		
18-24	-0.079	-0.011
	(-4.2)	
25-34	-0.059	-0.009
	(-4.0)	
35-54	-0.042	-0.006
	(-3.3)	
Tenure/100	-1.852	-0.279
	(-11.9)	
Tenure/100 Squared	4.559	0.687
	(10.2)	
Monthly Wage (predicted)	0.191	0.029
	(6.5)	
$\Phi(W_M)$ (predicted)	0.095	0.014
	(2.6)	
Constant	-0.541*	-0.081
	(-1.7)	

Table 4.2: Continued

Log-likelihood	-88476.8	
Chi-squared	33654.0	

Notes: (i) a set of industry, regional and time dummies are included in the specification;
(ii) AgeS5=(Age-5) if Age>5, 0 otherwise; AgeS10=(Age-10) if Age>10, 0 otherwise;
(iii) t-ratios are in parentheses;
(iv) all estimates are significant at 1%, except those with an *.

Table 4.3: Wage Equation - Two-step Tobit Results
Full-time Workers (N=266024)
Dependent variable: log of Real Monthly Wage

Variables	Coefficient	Marginal Effect
Female	-0.158	-0.152
	(-100.8)	
Education	0.050	0.048
	(160.6)	
Age/100	2.643	2.536
	(52.6)	
Age/100 Squared	-2.499	-2.397
	(-40.7)	
Tenure/100	0.479	0.460
	(15.0)	
Tenure/100 Squared	-0.492	-0.472
	(-5.3)	
Qualification Levels		
Non-defined	0.349	0.335
	(58.4)	
Manager and Highly Professional	0.742	0.712
	(101.5)	
Professional	0.542	0.520
	(84.6)	
Supervisors	0.421	0.404
	(77.0)	
Highly Skilled and Skilled	0.198	0.190
	(43.5)	
Semi-skilled and Unskilled	0.075	0.072
	(16.2)	
Firm Size	0.016	0.016
	(23.8)	
Sales per Worker	0.070	0.067
	(90.2)	
Regional Unemployment rate	-0.060	-0.058
	(-35.9)	
Probability of Displacement (predicted)	-0.479	-0.459
	(-37.0)	
Constant	9.356	8.976
	(678.1)	
Log-likelihood	-76453.8	
$\widehat{\sigma}$	0.31	

Notes: (i) a set of industry, regional and time dummies are included in the specification;
(ii) t-statistics are in parentheses;
(iii) all estimates are significant at 1%, except those with an *.

5 Conclusion

In this paper we have investigated how wages adjust to unfavorable shocks that raise the risk of displacement through firm closing, and to what extent a wage change affects the exit likelihood. For this purpose, a simultaneous-equations model was applied to a large longitudinally linked employer-employee data set of workers displaced due to firm closing. Three main conclusions emerge from this exercise.

First, after controlling for employers' heterogeneity and the skill composition of the workforce, the results indicated that wages have a positive impact on the failure rate. High-wage paying firms face higher exit rates than low-paying firms, *ceteris paribus*. Indeed, a 1% increase in monthly wages raises the probability of displacement through firm closing by 0.46%.

Second, a negative and strong impact of the probability of closing on wages was found, favoring the hypothesis that the risk of unemployment depresses wages. This result is consistent with the idea that pay is fixed in a bilateral bargain where unemployment acts to weaken workers' bargaining position [Blanchflower, (1991)]. Workers employed in firms at risk earn 3% less one year prior to closing than workers in firms with no risk of closure. This robust empirical evidence reinforces the hypothesis that instead of requiring a compensating differential for a higher risk of displacement, workers in firms at risk are able to agree upon wage concessions/moderation in order to avoid the firm's shutdown.

Third, minimum wage restrictions were seen to increase the failure rates. A high proportion of minimum wage earners in a firm may preclude the possibility of wage concessions in response to unfavorable shocks, and thus accelerate the exit decision. In other words, firms with a higher incidence of minimum wage earners are more vulnerable to adverse demand shocks due to their inability to adjust wages downward. In fact, beyond the direct effect of wages on the failure rate, a 1% increase in the incidence of minimum wage earners raises the probability of displacement by 0.028%.

References

[1] Addison, J., L. Bellmann and A. Köling (2002), "Unions, Works Councils and Plant Closings in Germany", IZA Discussion Paper N° 474.

[2] Audretsch, D. and T. Mahmood (1994), "The Rate of Hazard Confronting New Firms and Plants in U.S. Manufacturing", *Review of Industrial Organization*, Vol. 9, pp. 41-56.

[3] Bernard, A. and J. Jensen (2002), "The Death of Manufacturing Plants", NBER Working Paper N° 9026.

[4] Blanchard, O. and P. Portugal (2001), "What Hides Behind an Unemployment Rate: Comparing Portuguese and U.S. Labor Markets", *American Economic Review*, Vol. 91, pp. 187-207.

[5] Blanchflower, D. (1991), "Fear, Unemployment and Pay Flexibility", *Economic Journal*, Vol. 101, pp. 483-496.

[6] Capelli, P. (1985), "Plant-level Concession Bargaining", *Industrial and Labor Relations Review*, Vol. 38, pp. 90-104.

[7] Cooper, A., F. Gimeno-Gascon and C. Woo (1994), "Initial Human and Financial Capital as Predictors of New Venture Performance", *Journal of Business Venturing*, Vol. 9, pp. 371-395.

[8] Davis, S., J. Haltiwanger and S. Schuh (1996), *Job Creation and Destruction*, Cambridge: MIT Press.

[9] Doms, M. and J. Jensen (1998), "Comparing Wages, Skills, and Productivity between Domestically and Foreign-Owned Manufacturing Establishments in the United States" in R. Baldwin *et al.* eds., *Geography and Ownership as Bases for Economic Accounting*, NBER Studies in Income and Wealth, Vol. 59, Chicago and London: University of Chicago Press.

[10] Dunne, T. and M. Roberts (1990), "Wages and the Risk of Plant Closings", CES Working Paper N° 90-6.

[11] Dunne, T., M. Roberts and L. Samuelson (1989), "The Growth and Failure of U.S. Manufacturing Plants", *Quarterly Journal of Economics*, Vol. 104, pp. 671-698.

[12] Evans, D. (1987), "The Relationship Between Firm Growth, Size, and Age", *Journal of Industrial Economics*, Vol. 35, pp. 567-581.

[13] Hall, B. (1987), "The Relationship Between Firm Size and Firm Growth in the United-States Manufacturing Sector", *Journal of Industrial Economics*, Vol. 35, pp. 583-606.

[14] Hamermesh, D. (1988), "Plant Closings and the Value of the Firm", *Review of Economics and Statistics*, Vol. 70, pp. 580-586.

[15] Hamermesh, D. (1991), "Wage Concessions, Plant Shutdowns, and the Demand for Labor" in J. Addison ed., *Job Displacement: Consequences and Implications for Policy*, Detroit: Wayne State University Press, pp. 83-106.

[16] Hamermesh, D. (1993), *Labor Demand*, Princeton: Princeton University Press.

[17] Harris, R. and P. Hassaszadeh (2002), "The Impact of Ownership Changes and Age Effects on Plant Exits in UK Manufacturing, 1974-1995", *Economics Letters*, Vol. 75, pp. 309-317.

[18] Jacobson, L., R. LaLonde and D. Sullivan (1993), "Earnings Losses of Displaced Workers", *American Economic Review*, Vol. 83, pp. 685-709.

[19] Jovanovic, B. (1982), "Selection and the Evolution of Industry", *Econometrica*, Vol. 50, pp. 649-670.

[20] Katz, L. (1986), "Efficiency Wage Theories: a Partial Review", NBER Macroeconomics Annual, pp. 235-276.

[21] Kumar, M. (1985), "Growth, Acquisition Activity and Firm Size: Evidence from the United-Kingdom", *Journal of Industrial Economics*, Vol. 33, pp. 327-338.

[22] Machin, S. (1995), "Plant Closures and Unionization in British Establishments", *British Journal of Industrial Relations*, Vol. 33, pp. 55-68.

[23] Maddala, G. (1983), *Limited-Dependent and Qualitative Variables in Econometrics*, Cambridge: Cambridge University Press.

[24] Margolis, D. (1999), "Part-Year Employment, Slow Reemployment, and Earnings Losses: The Case of Worker Displacement in France", in J. Haltiwanger *et al.* eds., *The Creation and Analysis of Employer-Employee Matched Data - Contributions to Economic Analysis*, North-Holland: Elsevier Science B.V., pp. 375-416.

[25] Mata, J. and P. Portugal (1994), "Life Duration of New Firms", *Journal of Industrial Economics*, Vol. 42, pp. 227-245.

[26] Mata, J. and P. Portugal (2002), "The Survival of New Domestic and Foreign Owned Firms", *Strategic Management Journal*, Vol. 23, pp. 323-343.

[27] Mata, J., P. Portugal and P. Guimarães (1995), "The Survival of New Plants: Start-up Conditions and Post-Entry Evolution", *International Journal of Industrial Organization*, Vol. 13, pp. 459-481.

[28] Murphy, K. and R. Topel (1985), "Estimation and Inference in Two-Step Econometric Models", *Journal of Business and Economic Statistics*, Vol. 3, pp. 370-379.

[29] Smith, A. (1974), *The Wealth of Nations*, Penguin Books: edited by Andrew Skiner.

[30] Stevens, A. (1997), "Persistent Effects of Job Displacement: The Importance of Multiple Job Losses", *Journal of Labor Economics*, Vol. 15, pp. 165-188.

Annex
The Portuguese Labor Market: a "Snapshot"

1 Introduction

Spurred by a good performance over the last two decades, interest in the Portuguese labor market has widened in recent years. Two distinct features characterize the Portuguese labor market: on the one side, the existence of very rigid employment protection legislation and a sclerotic labor market, on the other side, a high aggregate wage flexibility.

According to OECD (1999), Portugal ranks high in all indicators of employment protection. In fact, rules governing the termination of permanent contracts are very strict and trigger large costs in case of dismissal. This stricter dismissal legislation is associated with lower turnover in the labor market, with both jobs and unemployment spells tending to last longer.[1] As pointed out by Blanchard and Portugal (2001), even though the unemployment rate in Portugal over the past 15 years is quite similar to the U.S. rate, the average unemployment duration in Portugal is more than three times that of the U.S. Symmetrically, flows of workers into unemployment are three times lower in Portugal. The image of Portuguese unemployment is that of a stagnant pool, with low flows in and out, and long unemployment duration.

The Portuguese labor market is characterized, within the European context, to exibhit low and strongly procyclical unemployment rates. Given the weaker labor mobility, the main reason that justifies this situation is the existence of a high aggregate wage flexibility. In fact, real wage behavior tends to closely follow the evolution of the unemployment rate, decreasing in periods of high unemployment and increasing in periods of low unemployment.

Indeed, some studies at the macro level have concluded for the existence of a high aggregate wage flexibility. In a study for the period 1983-92, Luz and Pinheiro (1993) found that wages respond to unemployment virtually without lags, whereas a three-quarter lag mediates the influence of past inflation on wage growth. Their conclusions argue in favor of the stability of both the Beveridge relationship and the Phillips-type wage equation over the period and lend support to the conventional view of wage flexibility in the Portuguese labor market with respect to contemporaneous (or virtually so) unemployment and inflation conditions. In a follow-up paper, Luz and Pinheiro (1994) proceeded to a comparative analysis of labor market behavior in Germany, France, United Kingdom, Spain and Portugal with data up to 1993, attempting to control

[1] In 2000, the incidence of long-term employment, which measures the proportion of employees that remain in the same job for 20 years or more, was 45% and the average unemployment duration 19.7 months [Banco de Portugal (2001)/*Annual Repport 2000*, Table II.4.7, p. 159].

for productivity growth by detrending the (log) real wages. They report a poor trade-off between detrended real wage and the unemployment rate for all countries except Portugal. Portugal was also the sole country in which both the Beveridge and wage curves presented no clear evidence of instability.

In sum, the Portuguese labor market may be characterized, on one side, by a high degree of employment rigidity, mainly associated with the firing costs legally imposed and, on the other side, by a high degree of wage flexibility (OECD, 1992).

The purpose of this Annex is to characterize briefly the main institutions that rule the Portuguese labor market. The summary provided here is drawn in large part from the work of Quaternaire Portugal and CETE (1996) and Bover *et al.* (2000).[2]

2 Labor Market Institutions

2.1 Wage Setting Institutions

2.1.1 Collective Bargaining

The Portuguese Constitution provides the juridical principles of collective bargaining and grants unions the right to negotiate. The effects of the agreements are formally recognized and considered valid sources of labor law.

Concerning the bargaining mechanisms, a distinction should be made between the *conventional* regime and the *mandatory* regime. Conventional bargaining results from direct negotiation between employers' and workers' representatives. A mandatory regime, on the other hand, does not result from direct bargaining between workers and employers, being instead dictated by the Ministry of Employment. The Ministry can extend an existing collective agreement to other workers initially not covered by it or it can create a new one, if it is not viable to extend the application of an existing document. A mandatory regime is applied when workers are not covered by unions, when one of the parties involved refuses to negotiate or bargaining is obstructed in any other way.[3] Therefore, the impact of collective bargaining goes far beyond union membership and the distinction between unionized and non-unionized workers or firms becomes meaningless.

Usually, collective negotiations are conducted at the industry or occupation level. Firm-level negotiation, which for a time was a common practice in large public enterprises, has lost importance. The law does not establish mechanisms of co-ordination between agreements reached in different negotiations; however

[2] For a detailed description see "Labour Market Studies - Portugal" by Quaternaire Portugal and CETE (1996). See also Bover *et al.* (2000), for a comparison between the Portuguese and Spanish labor markets.

[3] Beyond the existence of compulsive extension mechanisms, voluntary extensions are also possible, when one economic partner (workers' representative or employer) decides to subscribe to an agreement which it had initially not signed.

preference is given to vertical over horizontal agreements, and the principle of the most favorable condition to the worker generally applies.

Since most collective agreements are industry-wide, covering companies with very different sizes and economic conditions, their contents tend to be general, setting mimimum working conditions, in particular the base monthly wage for each category of workers, overtime pay and the normal duration of work.[4] Moreover, only a narrow set of topics is updated annually, and therefore the content of collective agreements is often pointed out as being too immobile and containing little innovation [see Dornelas (1999) and Leitão (1999)].

Whatever the wage floor agreed upon for each category of workers at the collective bargaining table, firms are free to pay higher wages, and they often deviate from that benchmark, adjusting to firm-specific conditions [see Aperta *et al.* (1994) and Portugal and Cardoso (2003)].

Freedom of union organization and union affiliation are recognized in Portuguese law. Currently, two main confederations share the leadership of workers' organization: the General Confederation of Portuguese Workers (CGTP), founded in 1975, and the Workers' General Union (UGT), founded in 1978. The two confederations are organized along political lines.

There are no good data about union membership in Portugal since trade unions are not obliged to report their membership. Even though, estimates have been reported by Stoleroff and Neumann (1993), Cerdeira (1997) and OECD (1994). Stoleroff and Neumann report rates of unionization at the end of the 1980s in the 34-40% range and, according to them, membership that is similar to that found in countries with a longer tradition of trade-unionism, such as Germany, Italy or the United Kingdom. Cerdeira reports that between 1974-78 the unionization rate averaged 53%, dropping during 1991-95 to 36%. According to OECD, trade union membership in Portugal in 1990 was 32%. Nevertheless, and due to the existence of extension mechanisms, the number of employees covered by the bargaining system is notably much higher.[5]

Employers are grouped into three national confederations defined according to economic industry bounds. There is a confederation of agriculture (CAP), another for wholesale and retail trade (CCP), and the most influential is the confederation of manufacturing (CIP).

The Portuguese system of industrial relations apparently presents features of a centralized wage bargaining system. Indeed, massive collective agreements, often covering a whole industry, predominate in the economy, while firm-level collective bargaining covers a low proportion (less than 10%) of the workforce. Moreover, trade union confederations, employers' federations and the Government meet at the national level each year to set a guideline for wage increases (the so-called "social concertation"). However, this guideline is not mandatory and merely orients the collective bargaining that follows. The Council for Social

[4] Dornelas (1999) points out that they seldom deal with health, hygiene and security in the workplace, or training and its impact on career progression, or social protection beyond the compulsory minimum.

[5] In 1993, and according to data from *Quadros de Pessoal*, 98.8% of the employees were covered by a collective bargaining mechanism [see Pinto *et al.* (1996), p. 250].

Concertation, latter replaced by the Social and Economic Council, was created in 1984 as a tripartite forum (government, workers and employers' representatives) with the aim of promoting "social concertation", but its role concerning income and wage policies remains limited.

On the other side, the fragmented nature of the trade union structure, the fragmented employers' associations and the multiplicity of bargaining units provides the system with a certain degree of decentralization. Even though collective bargaining in Portugal takes place at a sectoral level and most workers are covered by the bargaining system due to the existence of mandatory extensions, the coordination between bargaining units is rather limited. In fact, the right to negotiate is given upon every employer or employers' association and to every trade union (regardless of the number of affiliated members they represent), and the parties have the possibility of choosing the level of negotiation - regional, occupational, industrial or national. This leads to the existence of a diffuse and complex system of wage bargaining with negotiation fragmented and agreements multiplied.

2.1.2 Minimum Wages

A mandatory minimum monthly wage was set for the first time in Portugal in 1974, covering workers aged 20 or older and excluding agriculture and domestic servants. Currently, there is a legal minimum for private household workers and another one for general workers. Workers formally classified as apprentices receive just 80% of the full rate.

The minimum wage is updated annually by the parliament, under government proposal.[6] Decisions on the level of the minimum wage are taken on a discretionary basis, usually taking into account past and predicted inflation and after consulting the social partners.

According to the annual report of Banco de Portugal (1999), the ratio between minimum wages and average wages fell between 1986 and 1997. In 1986, the minimum monthly wage level as a proportion of the average monthly wage represented about 52%, whereas in 1997 this proportion fell to 42.5%. Data for 1995 indicate that the proportion of workers that received the minimum legal wage or a subminimum was about 13%, although this percentage has steadily declined since 1985.

2.2 Employment Protection

Concerning employment protection, there is a basic principle in the Portuguese law: workers are entitled to a severance pay whenever an employment contract termination occurs. In fact, in all cases of dismissals of workers with permanent contracts (except for disciplinary reasons) the worker is entitled to a payment

[6] The only exceptions are 1982, when it was not updated, and 1989, when it was updated twice.

equal to one month of pay for each year of service, subject to a minimum of three months pay.

Besides the pecuniary costs that employers have to bear, conditions in which a termination of a permanent contract is admissible are also regulated quite strictly. Indeed, individual dismissal is allowed in case of employee's culpable behavior, employee's failure to adapt to changes in the nature of his work, and lay-off. Lay-offs affecting more than one worker (firms with fifty workers or less) or more than four (firms with more than fifty workers) become, legally, a collective dismissal. In all cases of individual dismissal, written communication of impending dismissal to the employee and to the works council and/or union is required. This statement must give the reasons on which the dismissal is based. The worker and his representatives are then given the opportunity to dispute the employer's allegations. In case of economic lay-offs, the worker may also ask for Labor Inspectorate intervention in which case officials have to verify the validity of the arguments put forward by the employer. For all types of dismissal these procedures take at least three weeks.

60-days advance notification of a collective dismissal to the works council or union and to the Ministry of Employment is also required. Within 15-days consultations between the three parties are mandatory. Alternatives to lay-off, the number of dismissals, and ways to mitigate the effects of dismissal are all issues that must be addressed during this consultation process. Once an agreement is reached, each worker selected for dismissal must be notified of the impending job loss. This must be done at least 60 days before the date of dismissal. Otherwise, the worker is entitled to the corresponding pay. In short, the legislation on dismissals imposes a long, complex, and costly process on employers.

Fixed-term contracts, which were first regulated in 1976, were conceived as an instrument of flexibilization at the margin. In 1989, however, legislative changes made fixed-term contracts more restrictive while maintaining similar severance pay disbursements for temporary and permanent contracts.[7]

Fixed-term contracts are permitted for objective reasons: replacement of temporarily absent permanent workers; exceptional workload; seasonal work; and specific projects. They are also permitted for business start-ups, the launching of new activities of uncertain duration, and recruiting the long-term unemployed and those individuals searching for their first job. They have a minimum duration of six months unless they are justified on objective grounds, in which case no minimum applies, and a maximum duration of three years. Three successive renewals are the maximum allowed by law.

A severance bonus equal to two days pay for each complete month of the contract's elapsed duration is granted to the worker when the end of the contract is reached and not converted into a permanent one.[8]

[7] In 2000, fixed-term contracts represented 14% of total employment (Source: INE - *Inquérito ao Emprego 2000*).

[8] If the employer terminates the contract before its term, and the termination is unlawful, the worker is entitled to compensation equal to the pay loss from dismissal to the date of the court's decision or the term of the contract (whichever applies). He (or she) is also entitled

2.3 Unemployment Benefits

Before 1985 unemployment insurance was pratically nonexistant. In 1985 a regime of unemployment protection was introduced, establishing a true unemployment insurance as a part of the social security system. Therefore, in 1985, unemployment insurance benefits were introduced. At this time, the unemployment social assistance benefit was also modified. The amount of this benefit is independent of the individual's previous work background (means-tested social benefit). In 1989 eligibility criteria for unemployment benefits were eased and the maximum duration period increased, both for insurance and for social benefits. In 1995 the unemployment benefits coverage was 52% [OCDE (1998)].

To be eligible for unemployment insurance benefits, workers must meet two conditions. First, they must have been employed for at least 540 days during the previous 24 months. Second, they must be involuntarily unemployed and have registered as job seekers at the nearest Employment Office. The amount of benefit paid represents 65% of the average monthly earning during a 6-month period prior to the second month before becoming unemployed. This amount can neither be less than the legal minimum wage nor exceed by three times its value. The duration of these benefits varies from a maximum of 12 months for individuals under 25 years of age to a maximum of 36 months for those aged 55 or over.

All unemployed whose households have per capita income below 80% of minimum wage and do not qualify for regular unemployment benefits because those are already exhausted or because they were employed for insufficient time in the previous 2 years can have access to unemployment social assistance benefits. The amount received is calculated as a percentage of the minimum legal wage, which can vary from 70% to 100% depending on the number of dependents in the household. The maximum duration of this type of benefit is the same as for unemployment insurance benefits, unless these have already been exhausted. In that case, its duration is half of that mentioned above. For workers over 55, unemployment social assistance benefits are paid until the age of 60, when the individual can apply for early retirement.

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to reinstatement if the term of the contract has not yet been reached.

References

[1] Aperta, A., I. Moreira and M. Murteira (1994), *Análise das Diferenciações entre Remunerações Convencionais e Efectivas*, Ministério do Emprego e Segurança Social, Colecção Estudos, série B - 'Rendimentos'.

[2] Banco de Portugal (1999), Relatório Anual de 1998, Lisboa: Banco de Portugal.

[3] Banco de Portugal (2001), Annual Report 2000, Lisboa: Banco de Portugal.

[4] Blanchard, O. and P. Portugal (2001), "What Hides Behind an Unemployment Rate: Comparing Portuguese and U.S. Labor Markets", *American Economic Review*, Vol. 91, pp. 187-207.

[5] Bover, O., P. García-Perea and P. Portugal (2000), "Labor Market Outliers: Lessons from Portugal and Spain", *Economic Policy*, Vol. 31, pp. 381-428.

[6] Cerdeira, M. (1997), *A Evolução da Sindicalização Portuguesa de 1974 a 1995*, Ministério do Emprego e Segurança Social, Colecção Estudos, série C - 'Trabalho'.

[7] Dornelas, A. (1999), "As Relações Industriais em Portugal: É Possível Mudar? É Possível não Mudar?", *Sociedade e Trabalho*, Vol. 7, pp. 46-53.

[8] Leitão, M. (1999), "Traços Gerais da Contratação Colectiva em Portugal", *Sociedade e Trabalho*, Vol. 4, pp. 19-28.

[9] Luz, S. and M. Pinheiro (1993), "Desemprego, Vagas e Crescimento Salarial", Banco de Portugal, Boletim Trimestral 15-2.

[10] Luz, S. and M. Pinheiro (1994), "Wage Rigidity and Job Mismatch in Europe: Some Evidence", Banco de Portugal Working Paper 2-94.

[11] Organisation for Economic Cooperation and Development (1992), *OECD Economic Surveys, Portugal, 1991/92*. Paris: OCDE.

[12] Organisation for Economic Cooperation and Development (1998), *OECD Employment Outlook*. Paris: OCDE.

[13] Organisation for Economic Cooperation and Development (1999), *OECD Employment Outlook*. Paris: OCDE.

[14] Pinto, M., P. Martins and A. Carvalho (1996), *European Employment and Industrial Relations Glossary: Portugal*, London, Sweet and Maxwell, Luxembourg, Office for Official Publications of the European Communities.

[15] Portugal, P. and A. Cardoso (2003), "Bargained Wages, Wage Drift, and the Design of the Wage Setting System", *mimeo*.

[16] Quaternaire Portugal and Centro de Estudos de Economia Industrial, do Trabalho e da Empresa (1996), *Labour Market Studies: Portugal*, Employment & Social Affairs, European Comission.

[17] Stoleroff, A. and R. Naumann (1993), "A Sindicalização em Portugal: a Sua Medida e a Sua Distribuição", Sociologia-Problemas e Práticas, Vol. 14, pp. 19-47.