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Labor Contracts and Flexibility: Evidence from a Labor Market Reform in Spain

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Abstract

This paper evaluates the effects on employment, job turnover and productivity of a labor market reform in Spain that eliminated dismissal costs for fixed-term or temporary contracts. Our empirical results are based on a panel of 2356 Spanish manufacturing firms for the period 1982-1993. We postulate and estimate a dynamic labor demand model with indefinite and fixed-term labor contracts, and a general structure of labor adjustment costs. Experiments using the estimated model show important positive effects of the reform on total employment (i.e., a 3.5% increase) and job turnover. There is a strong substitution of permanent by temporary workers (i.e., a 10% decline in permanent employment). The effects on labor productivity and the value of a firm are very small. These effects contrast with the ones of a counterfactual reform consisting in halving firing costs of all type of contracts. That policy implies the same increase in total employment, but much larger improvements in productivity, and the value of firms.

Keywords: Labor demand, Firing costs, Temporary contracts, Estimation of dynamic structural models.

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1 Introduction

Regulation of workers' dismissal is one of the labor market institutions most commonly invoked to explain the large and persistent differences between European and North American unemployment rates. From a theoretical point of view, the effect of firing costs on employment is ambiguous. Firing costs reduce hiring during expansions, but these costs also reduce dismissals during downturns. The net effect depends on different factors, including the size of hiring and firing costs and the degree of persistence of demand and supply shocks. Therefore, the employment effect of firing costs is an empirical question that should be evaluated case by case. The labor market reforms that several European countries have implemented since the 1980s provide unique information to identify the effect of firing costs on firms' labor demand decisions. In this paper we study the effects on employment and firms' productivity of a Spanish labor market reform that took place in 1984, which eliminated previous restrictions to use temporary contracts and reduced firing costs under this type of contract.

The consequences of job security provisions on labor market performance have been broadly analyzed both at the theoretical and at the empirical level and using very different approaches. The studies differ in the data used (aggregate data, industry-level data, household and firm level data), the scope of the analysis (from the study of a particular country to cross-country comparisons), and on the methodological approach. The results are not conclusive: whereas some contributions provide evidence supporting that job security provisions have negative effects on employment and activity rates as well as on the speed of adjustment of employment and output, some others find negligible effects on the level of employment.

The main purpose of this paper is to evaluate the effects on employment, job turnover and productivity of a labor market reform in Spain occurred in 1984, which allowed the widespread use of fixed-term or temporary contracts and reduced the redundancy payments at termination of these contracts. Before the reform, the use of temporary contract was subject to the principle of causality, so that these contracts had already been used to a certain extent in agriculture, construction and services industries; however, their use in

manufacturing had been very scarce. After the reform, temporary contracts could be applied by any firm, irrespective of their size, industry or performance, to any type of worker, irrespective of their occupation, age, or sex. Nevertheless, the stringent dismissal regulations for indefinite-duration or permanent contracts remained unchanged after the reform.

Our approach in this paper combines the estimation of a micro-econometric dynamic structural model of labor demand with a comparison of estimated structural parameters before and after the policy change in 1984. Our data set just covers 2,356 Spanish manufacturing firms during the period 1982-1993, so that we exploit the time variation in the policy rule, before and after 1984. We estimate a dynamic structural model of labor demand for permanent and temporary employment both before and after the reform. In the context of our model, the reform implies a change in firing and hiring costs of temporary workers as well as in the relative productivity of this type of labor. Our estimates show a significant reduction after the reform in the costs of hiring and firing temporary workers. Experiments using the estimated model show important positive effects of the reform on total employment (i.e., a 3.5% increase) and job turnover. There is a strong substitution of permanent by temporary workers (i.e., a 10% decline in permanent employment). The effects on labor productivity (and the value of a firm are very small. These effects contrast with the ones of a counterfactual reform consisting in halving firing costs of all type of contracts. That policy implies the same increase in total employment, but much larger improvements in productivity, and the value of firms.

However, its effects on the value of a firm are negligible, and the effects on productivity are negative. These effects contrasts with the sizeable increases in output and value of firms under a hypothetical reduction in firing costs for all type of contracts. Compared with this counterfactual reform, the factual introduction of temporary contracts leads to excess turnover and employment of workers with low firm-specific experience.

Our approach allows us to overcome some important limitations of an empirical strategy based on reduced form evidence. First, the fact that the reform under study was applicable

to any type of firm and to any type of worker, makes implausible a differences-in-differences approach based on comparing the outcomes before and after the reform between agents affected and unaffected by such reform. Second, given that some other institutional changes took place in Spain after 1984 (for instance, the Spanish entry in the European Economic Community in 1986), a reduced form approach does not ensure that we are controlling for the sort of structural changes that we want to consider, that is, those which affected firing costs of temporary workers. And third, we are also interested in the evaluation of counterfactual policies.

There is a large literature on the structural estimation of dynamic structural models of labor demand that goes back to the seminal paper by Sargent (1978).[1] Our model builds on and extends recent papers on dynamic structural models of labor demand with non-convex adjustment costs, such as Rota (2004), and Cooper and Willis (2004). The most relevant extensions are the following: (i) we consider two types of labor contracts, temporary and permanent; (ii) our specification of labor adjustment costs is very general and allows for fixed, linear and quadratic adjustment costs which can be different for the two types of contracts; and (iii) the specification of the unobserved variables in the econometric model is quite flexible and it includes unobservables in the production function, in the marginal costs, and in the fixed costs of the two types of labor.

The rest of the paper is organized as follows. In section 2, we provide an overview of the previous literature on the effects of job security provisions, and describe the institutional features of the Spanish labor market. Section 3 describes our dataset. In section 4, we explain the theoretical model as well as our key identification assumptions to evaluate the effects of the policy change. The estimation results of the structural model are provided in section 4. We present experiments that evaluate the effects of the reform in section 5. Section 6 summarizes our main findings and concludes.

[1] See Hamermesh (1993) and the recent survey paper by Bond and Van Reenen (2007) for references.

2 The role of job security provisions

2.1 Previous evidence

There is a broad and growing literature on the consequences of job security provisions on the labor market. A first line of research uses longitudinal data of countries in order to evaluate the effects of severance pay on several labor market outcomes exploiting the differences across countries. Using a panel of OECD countries and constructing two alternative measures of severance pay, Lazear (1990) found that severance pay has negative effects on employment and activity rates, and a positive effect on unemployment. Addison and Grosso (1996) corroborate the positive influence of severance pay on unemployment, but they find very little evidence "to suggest that its contribution to rising unemployment is material". Burgess, Knetter and Michelacci (2000) evaluated the effects of job security provisions on the adjustment speed of employment and output, using longitudinal data on the seven largest OECD countries disaggregated by 2-digit industries, which allows to control for differences in adjustment speed among industries. Their results point out that less regulated countries show a faster adjustment. Using a sample of OECD and Latin American countries, Heckman and Pagés (2004) found a strongly negative effect of job security provisions on employment rates, such effect varying substantially among different types of workers.

A similar line of research has evaluated the consequences of job security regulations by means of comparing a small number of countries. Abraham and Houseman (1993, 1994) compare the adjustment speed of employment and hours in manufacturing industries in response to demand shocks in several European countries (Germany, France and Belgium) and in the United States. Their main finding is that the higher costs of adjusting employment levels in European countries are compensated by the lower costs of adjusting average hours, and therefore there are no substantial differences in the adjustment of total labor input. Bover, García-Perea and Portugal (2000) try to explain why unemployment rates in Spain and Portugal are so different even though their labor market institutions appear to share many similarities. The primary factor explaining the much higher unemployment rate in

Spain appears to be its lower level of wage flexibility in Spain, combined with a much more generous system of unemployment insurance.

A second line of research has addressed the effects of severance payments on employment by means of calibration of theoretical models. Bentolila and Bertola (1990) calibrate a partial equilibrium labor demand model using aggregate data of several European countries, obtaining negligible effects. In a similar setting, Bertola (1990) finds that job security provisions do not necessarily lower average employment unless further restrictions on wage flexibility, such as minimum wage legislation, operate. In contrast, Hopenhayn and Rogerson (1993) calibrate a general equilibrium model using US firm-level data and considering entry and exit of firms, obtaining that an introduction of firing costs would reduce employment substantially. Cabrales and Hopenhayn (1997) calibrate a similar model using firm-level evidence on job matches before and after the 1984 reform which allowed the widespread use of temporary contracts, finding that the reform has induced a large increase in the turnover rate but a moderate effect on employment. Güell (2000) analyzes the effects of temporary contracts in the context of an efficiency wage model, concluding that the introduction of this type of contract need not to increase aggregate employment.

A third line of research has exploited data before and after specific reforms in the labor market in order to evaluate how changes in job security provisions has affected labor market outcomes using a differences-in-differences approach. Kugler (2004) studies the effect of a reduction in firing costs in Colombia with data before and after such reform. The fact that the Colombian labor market is broken down in workers covered and non covered by the legislation provides a treatment and a control group which permits a differences-in-differences approach by comparison between the unemployment hazards for these two groups before and after the reform. The results provide evidence about the negative effect of severance payments on employment. Hunt (2000) exploited industry-level German data to conclude that the German reform in 1985 which facilitated the use of temporary contracts did not affect employment adjustment. In a similar line but with a different approach, Bentolila

and Saint-Paul (1992) use firm-level data to evaluate the effect of a Spanish reform which introduced temporary contracts and find a rise in the speed of adjustment, although they do not use data before the reform.

Our approach in this paper combines the estimation of a dynamic structural model of labor demand with a comparison of estimated structural parameters before and after the policy change in 1984. There is a large literature on the structural estimation of dynamic structural models of labor demand that goes back to the seminal paper by Sargent (1978). As mentioned in the Introduction, our model builds on and extends the recent literature on dynamic structural models of labor demand with non-convex adjustment costs (Rota, 2004, and Cooper and Willis, 2004). As far as we know, this is the only dynamic structural model of labor demand that has been used to evaluate a labor market reform.

2.2 Labor contract regulations in Spain

According to the OECD, the Spanish labor market is among the most regulated in Europe. Job security rules and, in particular, strong mandatory severance payments, contribute importantly to the rigidity of such regulations.[2] The 1984 reform, which eliminated most of the previous restrictions to the use of temporary contracts, has been one of the major legal changes of the Spanish labor market in the last two decades. To understand the motivation of this reform and the context in which it took place, we provide a description of the institutional background before the reform, and the subsequent changes occurred.

During Franco's regime (1939-1975), the Spanish labor market was characterized by a hyper-regulated system of industrial relations under the monitoring of a single "union" to which both employers and employees had to belong. The prohibition of trade unions and the practical absence of collective bargaining were "compensated" with regulations that guaranteed full employment stability: in practice, most jobs were full-time jobs of indefinite

[2] Bentolila and Dolado (1994), using OECD data for selected European countries, find striking differences in regulations about authorization procedures for dismissals and mandatory severance payments for fair and unfair dismissals, with Denmark and the UK having the less severe, and France, Greece, Portugal and Spain being the countries with the most stringent regulations.

duration. This institutional background was transformed progressively after Franco's death in 1975. The first important change came in 1977 with the Royal Decree of Industrial Relations. The official single union was dismantled and free trade unions were legalized. Although the decree also recognized new grounds for fair dismissals based on economic reasons and simplified the legal procedures for collective redundancies, job security rules were basically unchanged.

In 1980, the *Estatuto de los Trabajadores* (Workers' Statute, ET hereinafter) established the conditions for a modern system of collective bargaining comparable to the ones prevailing in other democratic European countries. However, it maintained many of the legal and administrative restrictions on dismissals. For *permanent* workers (those with an indefinite-term contract), mandatory severance payments were 20 days of salary per year of job tenure (up to a maximum of 1 year wages) if the dismissal is considered 'fair', and 45 days (up to a maximum of 42 months wages) if it is considered 'unfair'. In principle, there are two types of fair reasons: those attributable to the worker, when he is considered incompetent or negligent to perform the tasks for which he was hired, and objective reasons that cannot be attributed to the worker (for economic or technological reasons). However, the scope of the second reason was very limited. Furthermore, the *burden of proof* for a fair dismissal must be assumed by the firm (see Bentolila, 1997). If the worker does not accept the dismissal –as it is usually the case–, he may sue the firm for unfair dismissal. This obliges the firm to undertake a legal process to prove the fairness of the dismissal, and during this process the firm should assume the legal costs in any case, as well as the salaries of the worker (*procedure wages*) in the case that the dismissal is legally declared unfair. Given that the labor courts are in many cases favorable to the workers, the agreed severance payments can even exceed the statutory amounts for unfair dismissals. Another legal requirement is the mandatory advance notice of 30 days. These job security rules for *permanent* workers have remained unchanged until 1997.[3]

[3] In 1997, trade unions and employer organizations signed the *Acuerdo Interconfederal para la Estabilidad del Empleo* (National Agreement on Employment Stability). This agreement led to a new permanent contract

The ET provided the possibility of *temporary* or fixed-term contracts, which could be cancelled at termination with a much smaller severance payment and without court or regulatory intervention. However, the use of temporary contracts was mainly limited to jobs that were temporary in nature because of the seasonal nature of the production activity, the need to cover absent posts, or the start-up of a new firm.

Figure 1 shows the evolution of GDP growth and unemployment in Spain. Despite the fact that GDP was growing at the beginning of the eighties, the unemployment rate followed its increase, and by the end of 1984, unemployment in Spain was close to its peak (about 21%), and the Spanish economy was suffering the dismantling process of obsolete plants in the heavy industries. This fact, together with the complaints of entrepreneurs about the rigid employment legislation, forced the government to broaden the scope of temporary contracts in an attempt to boost employment. The ET was reformed in 1984, introducing the most important legal change of the Spanish labor market in the previous two decades, removing most of the restrictions on non-causal fixed-term contracts. The main feature of the reform is that the use of temporary contracts is no longer linked to the principle of causality, so that they could be applied to any activity, temporary or not, and to any type of firm or worker. Furthermore, they might be signed for short periods (three, six or twelve months), firing costs at termination were low (12 days of wages per year of tenure) or even zero in some cases, and their extinction could not be appealed to labor courts. Nevertheless, an important limitation for the use of temporary contracts under the new law was that they could be renewed only up to three years. After this period the firm should decide whether to offer the worker a permanent contract or to dismiss him.[4] Importantly, the reform did not alter the stringent dismissal regulations for permanent or indefinite-duration contracts.

After this reform, the number and the proportion of temporary jobs in the Spanish economy increased sharply. In Figure 2 we present the evolution of the proportion of temporary

which maintained the severance payments for fair dismissals but lowered those for unfair dismissals to 33 days of salary (up to 42 months wages), yet their utilization was limited to certain type of workers.

[4] Furthermore, if a firm lays off a temporary worker, it must wait for a year in order to hire him again.

employment in total employment since 1987.[5] The share of temporary contracts in total employment, which was estimated to be about 10% of total employment and 3% of manufacturing employment in 1984, rose to 35 and 30 percent, respectively, in 1995, and have remained at high levels since then.[6] Spain has become by far the European country with the highest percentage of temporary employment, with temporary contracts representing 80% of hires in the period 1986-1990 (Bentolila and Dolado, 1994). This important increase in temporary employment points out that firms have found these contracts attractive to reduce firing costs. Nevertheless, this behavior is consistent with either positive or negative employment effects of the reform. Evaluating the effects of the reform on employment and output requires to analyze how individual firms' hiring and firing decisions have changed after the reform.

3 Data and preliminary evidence

The main data set has been taken from the database of the Balance Sheets of the Bank of Spain (CBBE hereafter), which contains firm-level annual information on the balance sheets and other complementary information on economic variables, such as employment by type of contract, output, physical capital and the total wage bill. The sample consists on an unbalanced panel of non-energy manufacturing firms with a public share lower than 50 percent for the period 1982 to 1993. To obtain the final sample of 2,356 firms we have eliminated those for which some of the following variables were negative or took implausible values: book value of capital stock, sales, gross output, total labor costs, permanent employment, and temporary employment. Due to the fact that response is completely voluntary, largest firms are over-represented in the sample. The firms included in this sample represent 40%

[5] Unfortunately, the Spanish Labor Force Survey did not reported any information about the type of contract before 1987. In section 3, we present descriptive evidence on the evolution of temporary employment for the period 1982-1993 using our panel of manufacturing firms from the CBBE database.

[6] Figure 2 shows a large disparity between the proportion of temporary worker from the Labor Force Survey and the proportion from the CBBE. The main factor to explain this discrepancy is that CBBE over-represents large manufacturing firms, and this type of firms tend to have a smaller proportion of temporary workers. See Figure 4 below.

of total Spanish manufacturing value added during the period. Table 1 presents the sample distribution of firms by industry and size.

The CBBE contains firm-level information on the number of workers by type of contract (temporary or permanent), and on the average duration (in weeks) of temporary contracts. To maintain measurement consistency, number of temporary employees is calculated in annual terms by multiplying the number of temporary employees along the year times the average number of weeks worked by temporary employees and divided by 52. It is worth to notice that, as it happens in most firm-level datasets, there is not information on employment flows along the year, and therefore only net employment changes are observed. Gross output at retail prices is calculated as total sales, plus the change in finished product inventories and other income from the production process, minus taxes derived on the production (net of subsidies). Real output has been obtained using as deflator the Retail Price Index at 3 digits industry level. The information on the firm's total wage bill (which allows to calculate the average wage rate for total employees at the firm-level) is not broken down by type of contract. Wage information by type of contract is available from the Spanish Wage Distribution Survey (*Distribución Salarial*, DS hereinafter). However, the DS dataset is only available for the years 1988 and 1992, and it provides only aggregate information. We describe in Section 5.1 our approach to obtain estimates of wages of permanent and temporary workers for the whole period 1982-1993.

In Figure 3, we report the evolution of the growth rates of real output and employment for our sample of firms. The evolution of real output growth shows that the period 1982-1993 covers an expansion, 1986-1989, and a recession 1990-1993. However, the number and the proportion of permanent employees have monotonically decreased along the sample period, as shown in Figure 2. After the introduction of the new regulation of temporary contracts in November 1984, temporary employment rose significantly from 1986 to 1990 and decreased during the economic downturn from 1990 to 1993, and its share in total employment rose from 2.89 percent in 1985 to 9.72 percent in 1993. Although the evolution

of temporary employment in our sample keeps coherency with the aggregate series for the overall economy, and particularly with the aggregate series for the manufacturing industry (in fact, the correlation coefficient between both series is above 90 percent), the figure for our sample is clearly much smaller than the aggregates figures, which were well above 20 percent at the beginning of the nineties. This discrepancy is due to the fact that larger companies, which are over-represented in our sample, are more prompted to use permanent employment than small or medium ones. In Figure 4, we can see that the proportion of temporary employment for firms in our sample differs very much between large firms, for which this proportion is lower, and small and medium firms.

Figure 5 presents the job creation and job destruction rates for permanent and temporary employment using the statistics proposed by Davis and Haltiwanger (1992).[7] The small job turnover rates for permanent employment contrasts with the very high rates for temporary employment. Furthermore, the creation and destruction rates for temporary employment are much more correlated with the cycle than those for permanent employment. This is evidence of how firing costs can have very important effects on job turnover rates. It also reflects the fact that although the reform introduced larger flexibility for new hires, it kept the core of permanent employees unaffected.

Figure 6 presents the times series of the proportion of firms with positive, negative and zero annual change in permanent employment. We observe a remarkable frequency of no adjustments in permanent employment (about 19%), which is fairly stable over the cycle, suggesting an important persistence in permanent employment. This evidence is consistent with the existence of lump-sum or kinked adjustment costs.

Table 2 presents descriptive statistics on employment and productivity for a pre-reform period (1982-1984) and a post-reform period (1989-1992). For the sake of comparability, we consider a common subsample of firms in the two periods (389 firms). The post-reform

[7] Our measures are based on firm level data instead of plant level data as in Davis, Haltiwanger (1992) for US. This can be a factor, in addition to the different labor market institutions in Spain and US, that contributes to the smaller job turnover rates that we find in our data.

period 1989-1992 has been also selected for comparability reasons: i.e., as shown in the first row of Table 2, the cross-sectional distributions of the rates of growth in real output are very similar in 1982-1984 and 1989-1992. From this comparison we can establish some interesting facts. The cross-sectional distribution of employment growth has the same median in the two periods, but it is significantly more disperse after the reform. The proportion of temporary employment increases, and as result of this there is a small increase in total employment despite a reduction in permanent employment takes place. Interestingly, we can see that firm productivity, as measured by the sales to wage bill ratio, goes down after the reform. Of course, these changes could be, or not, consequence of the labor market reform. To obtain more robust measures of the effects of the reform, we estimate a structural model.

4 Model

4.1 Basic assumptions

Consider an economy with two types of labor contracts: fixed-term and indefinite-duration contracts. We denote employees as temporary or permanent depending on whether they enjoy a fixed-term or an indefinite-duration contract, respectively. We assume that a fixed-term contract lasts only one period (year). In principle, the only exogenous feature that distinguishes a permanent and a temporary contract lies in the dismissal costs. Firms are enforced by law to pay a severance to each dismissed permanent worker, but temporary workers are not entitled to any compensation upon dismissal. Although dismissal costs appear as the only exogenous difference between these two contract types, they can generate, endogenously, further differences between workers. Particularly, two major differences are expected to appear. On the one hand, incentives to invest in firm-specific human capital are stronger for workers with indefinite-term contracts than for those with fixed-term contracts. This fact might create a productivity gap between permanent and temporary workers. On the other hand, the higher costs of dismissals will place permanent workers in a better bargaining position within the firm. This fact might induce a wage gap between permanent

and temporary workers. We incorporate these differential features in our model, yet we take them as exogenous for the sake of simplicity.

Firms produce a homogeneous good using labor as the only variable input, and sell their output in a competitive market.[8] Every period t, the firm chooses the amounts of permanent and temporary labor that maximize its expected intertemporal profit, $E_t\left(\sum_{j=0}^{\infty}\beta^j\Pi_{t+j}\right)$, where E_t is the conditional expectation function given the information up to period t, Π_t denotes profits at period t, and $\beta \in (0,1)$ is the discount factor. Current profits, measured in output units, are:

$$\Pi_t = Y_t - WB_t - AC_t + \xi_t, \tag{1}$$

where Y_t is real output, WB_t is the wage bill, AC_t represents labor adjustment costs, and the term ξ_t contains other components of current profit which are observable to the firm but unobservable to the econometrician. Physical capital is treated as a component of the firm idiosyncratic shock and it is assumed to follow an exogenous process.

The production technology is described by the production function

$$Y_t = (L_t^P + \lambda\, L_t^T)^{\alpha_L}\ \exp(\eta_t), \tag{2}$$

where L_t^P and L_t^T represent the corresponding amounts of firm's permanent and temporary workers; $\alpha_L \in (0,1)$ and $\lambda \in (0,1)$ are parameters; and η_t is an exogenous and idiosyncratic productivity shock. The parameter λ measures the productivity of temporary workers with respect to permanent workers. The productivity shock is assumed to follow a first-order Markov process with transition probability function $f_\eta(\eta_{t+1}|\eta_t)$.

The wage bill is $WB_t = W_t^T L_t^T + W_t^P L_t^P$, where W_t^T and W_t^P are the wages of temporary and permanent workers, respectively. The wage of temporary workers is determined at the market level, and it is the same for all firms operating in the market. However, the wage of permanent workers is firm-specific (e.g., internal labor market, rent-sharing). The pair of

[8] Alternatively, we may consider that firms compete in monopolistic product markets with isoelastic demand curves. In that setting, our production function should be re-interpreted as a revenue function, and its parameters are a combination of technological parameters and the elasticity of demand.

wages $W_t = (W_t^T, W_t^P)$ follows a first-order Markov process with transition probability function $f_W(W_{t+1}|W_t)$. Section 5.1 presents our specification assumptions on the joint dynamics of the wages of permanent and temporary workers.

The specification of labor adjustment costs, defined in terms of net employment changes, includes both lump-sum and linear components:

$$\begin{aligned} AC_t &= 1(\Delta L_t^T > 0)\ \left(\theta_{H0}^T + \theta_{H1}^T \Delta L_t^T\right) + 1(\Delta L_t^T < 0)\ \left(\theta_{F0}^T - \theta_{F1}^T \Delta L_t^T\right) \\ &+ 1(\Delta L_t^P > 0)\ \left(\theta_{H0}^P + \theta_{H1}^P \Delta L_t^P\right) + 1(\Delta L_t^P < 0)\ \left(\theta_{F0}^P - \theta_{F1}^P \Delta L_t^P\right) \end{aligned} \tag{3}$$

where 1(.) is the binary indicator function; $\Delta L_t^j \equiv L_t^j - L_{t-1}^j$ denotes the net change in employment of type j; and $\{\theta_{H0}^T,\ \theta_{H1}^T,\ \theta_{F0}^T,\ \theta_{F1}^T,\ \theta_{H0}^P,\ \theta_{H1}^P,\ \theta_{F0}^P,\ \theta_{F1}^P\}$ are (non-negative) parameters. The first two summands refer to hiring and firing costs of temporary workers. The third and fourth terms are the adjustment costs for hiring and firing permanent workers, respectively.

In this model, workers within the same firm and with the same contract type are identical. Therefore, a firm can hire or fire permanent (temporary) workers, but it is never optimal to hire and fire simultaneously workers with the same type of contract. Of course, it can be optimal to hire permanent (temporary) workers and fire simultaneously temporary (permanent) ones. This said, it is straightforward to see that optimal decisions on employment can be expressed in terms of net employment change.

The firm chooses employment changes so as to maximize its expected intertemporal profit. We consider a discrete choice model such that the set of possible values of $(\Delta L_t^P, \Delta L_t^T)$ is discrete and finite. The main reason why we consider a discrete model is that there is much lumpiness in these employment decisions. In our data, the frequency of zeroes in annual employment changes is 18.8% for permanent employment and 49.1% for temporary employment. Furthermore, the frequency of employment changes within -5 and +5 workers is 65.5% for permanent labor and 80.4% for temporary labor. Let D be the finite set of possible discrete values for $(\Delta L_t^P, \Delta L_t^T)$.

The component of current profits that is unobservable from the point of view of the

econometrician is specified as follows:

$$\xi_t \equiv \xi(\Delta L_t^P, \Delta L_t^T, \varepsilon_t) = \sigma_P \ \varepsilon_t^P \ \Delta L_t^P + \sigma_T \ \varepsilon_t^T \ \Delta L_t^T + \sigma_0 \ \varepsilon_t^0(\Delta L_t^P, \Delta L_t^T) \tag{4}$$

σ_P, σ_T and σ_0 are parameters. We use $\sigma_{\boldsymbol{\varepsilon}}$ to represent the vector $(\sigma_P, \sigma_T, \sigma_0)'$. ε_t^P and ε_t^T are mutually independent standard normal random variables which are independently distributed over time and across firms. For every possible pair of discrete values $(j,k) \in D$, the variable $\varepsilon_t^0(j,k)$ is a logit error which is independently and identically distributed over time and across firms with type I Extreme Value distribution. We use $\boldsymbol{\varepsilon}_t$ to denote the vector of unobservables $\{\varepsilon_t^P, \varepsilon_t^T, \varepsilon_t^0(k,j) : (k,j) \in D\}$. This combination of normal errors and logit errors resembles the specification in the mixed (or random coefficients) multinomial logit model in McFadden and Train (2000). The existence of the terms $\varepsilon_t^P \Delta L_t^P$ and $\varepsilon_t^T \Delta L_t^T$ imply that unobservables are correlated across choice alternatives. That is, the model does not have the so called Independence of Irrelevant Alternatives property, which is a well known limitation of the standard multinomial logit model. Furthermore, the model retains some features of the multinomial logit which are convenient for estimation and to ensure good properties of the maximum likelihood estimator.

Every period, the firm has perfect knowledge about its stocks of permanent and temporary labor, wages, and the realized values of productivity and cost shocks, but it has uncertainty about the future values of these shocks. Let $\mathbf{x}_t \equiv (L_{t-1}^P, L_{t-1}^T, W_t, \eta_t)$ be the vector of state variables at period t, excluding $\boldsymbol{\varepsilon}_\mathbf{t}$. And let d_t be an categorical variables that represents the pair of decision variables $(\Delta L_t^P, \Delta L_t^T)$. The profit function can be written as:

$$\Pi_t = z\left(d_t, \mathbf{x}_t\right) \ \boldsymbol{\theta} + \xi(d_t, \boldsymbol{\varepsilon}_t) \tag{5}$$

where $\boldsymbol{\theta}$ is the 9×1 vector $\left(1, \theta_{H0}^T, \theta_{H1}^T, \theta_{F0}^T, \theta_{F1}^T, \theta_{H0}^P, \theta_{H1}^P, \theta_{F0}^P, \theta_{F1}^P\right)$, and $z\left(d_t, \mathbf{x}_t\right)$ is a 1×9 vector with corresponding elements

$$\begin{pmatrix} Y_t - WB_t \ , \ 1(\Delta L_t^T > 0) \ , \ 1(\Delta L_t^T > 0)\Delta L_t^T \ , \ 1(\Delta L_t^T < 0) \ , \ -1(\Delta L_t^T < 0)\Delta L_t^T \\ , \ 1(\Delta L_t^P > 0) \ , \ 1(\Delta L_t^P > 0)\Delta L_t^P \ , \ 1(\Delta L_t^P < 0) \ , \ -1(\Delta L_t^P < 0)\Delta L_t^P \end{pmatrix} \tag{6}$$

We can represent the firm's decision problem using the Bellman equation:

$$V(\mathbf{x}_t, \boldsymbol{\varepsilon}_t) = \max_{d_t \in D} \left[z\left(d_t, \mathbf{x}_t\right) \boldsymbol{\theta} + \xi(d_t, \boldsymbol{\varepsilon}_t) + \beta \sum_{x_{t+1}} \int V(\mathbf{x}_{t+1}, \boldsymbol{\varepsilon}_{t+1}) \; f_\varepsilon(d\boldsymbol{\varepsilon}_{t+1}) \; f_x(\mathbf{x}_{t+1}|\mathbf{x}_t, d_t) \right] \tag{7}$$

where $f_{\boldsymbol{\varepsilon}}$ is the density function of $\boldsymbol{\varepsilon}_{\mathbf{t}}$, and the transition of the vector of state variables $\mathbf{x}_t$ is:

$$f_x(\mathbf{x}_{t+1}|\mathbf{x}_t, d_t) = 1\{\mathbf{L}_t = \mathbf{L}_{t-1} + \boldsymbol{\Delta}\mathbf{L}_t\} \; f_\eta(\eta_{t+1}|\eta_t) \; f_W(W_{t+1}|W_t) \tag{8}$$

4.2 Probabilistic representation of firms' employment decisions

In this section, we follow Aguirregabiria and Mira (2002) to represent the optimal solution of the previous dynamic labor demand model as the unique fixed point of a mapping in probabilistic space. Let X be the space of possible values of $\mathbf{x}_t$. A firm's optimal behavior can be represented using a vector of *optimal choice probabilities* $\mathbf{P} = \{P(d|\mathbf{x}) : (d, \mathbf{x}) \in D \times X\}$, where $P(d|\mathbf{x})$ is the probability that the optimal decision at period t is d conditional on the value of $\mathbf{x}_t$ being $\mathbf{x}$. Define also $\boldsymbol{\pi}$ as the vector of parameters that characterize the transition probabilities f_η and f_W. Following Aguirregabiria and Mira (2002), we can obtain the vector $\mathbf{P}$ as the unique fixed point of a contraction mapping in the space of conditional choice probabilities: $\mathbf{P} = \Psi_{\theta,\sigma_\varepsilon,\boldsymbol{\pi}}(\mathbf{P})$, where $\Psi_{\theta,\sigma_\varepsilon,\boldsymbol{\pi}}(.)$ is the mapping. Under our assumptions on the probability distribution of $\varepsilon_{\mathbf{t}}$, this mapping is:

$$\Psi_{\theta,\sigma_\varepsilon,\boldsymbol{\pi}}(\mathbf{P})(d|\mathbf{x}) =$$

$$\int \frac{\exp\left\{ Z_{\boldsymbol{\pi},\mathbf{P}}\left(d, \mathbf{x}\right) \frac{\theta}{\sigma_0} + e_{\boldsymbol{\pi},\mathbf{P}}\left(d, \mathbf{x}\right) + \varepsilon^P \Delta L^P(d) \frac{\sigma_P}{\sigma_0} + \varepsilon^T \Delta L^T(d) \frac{\sigma_T}{\sigma_0} \right\}}{\sum_{j \in D} \exp\left\{ Z_{\boldsymbol{\pi},\mathbf{P}}\left(j, \mathbf{x}\right) \frac{\theta}{\sigma_0} + e_{\boldsymbol{\pi},\mathbf{P}}\left(j, \mathbf{x}\right) + \varepsilon^P \Delta L^P(j) \frac{\sigma_P}{\sigma_0} + \varepsilon^T \Delta L^T(j) \frac{\sigma_T}{\sigma_0} \right\}} \phi(d\varepsilon^P)\phi(d\varepsilon^T) \tag{9}$$

ϕ is the pdf of the standard normal. $\Delta L^P(d)$ and $\Delta L^T(d)$ represent the values for the change in permanent and in temporary employment, respectively, under choice alternative d. $Z_{\boldsymbol{\pi},\mathbf{P}}\left(d, \mathbf{x}\right)$ is a 1×9 vector with the present value of the stream of current and future values of the vector $z\left(.,.\right)$ conditional on the current value of $(d_t, \mathbf{x}_t)$ being $(d, \mathbf{x})$ and under

the assumption that the firm will behave in the future according to the choice probabilities in $\mathbf{P}$. For instance, the first element of the vector $Z_{\boldsymbol{\pi},\mathbf{P}}(d,\mathbf{x})$ is the present value of output minus wage bill. $e_{\boldsymbol{\pi},\mathbf{P}}(d,\mathbf{x})$ is a scalar with a similar interpretation as $Z_{\boldsymbol{\pi},\mathbf{P}}(d,\mathbf{x})$ but it contains the present value of the stream of future realizations of $\xi(d_{t+j},\boldsymbol{\varepsilon}_{t+j})$ associated with optimal future choices. More formally, we have that:

$$Z_{\boldsymbol{\pi},\mathbf{P}}(d,\mathbf{x}) \;=\; z(d,\mathbf{x}) + \textstyle\sum_{j=1}^{\infty}\beta^j\; E\left(z\left(d^*_{t+j},\mathbf{x}_{t+j}\right)\;|\; d_t=d,\mathbf{x}_t=\mathbf{x};\boldsymbol{\pi},\mathbf{P}\right) \tag{10}$$

and

$$e_{\boldsymbol{\pi},\mathbf{P}}(d,\mathbf{x}) \;=\; \textstyle\sum_{j=1}^{\infty}\beta^j\; E\left(\xi(d^*_{t+j},\boldsymbol{\varepsilon}_{t+j})\;|\; d_t=d,\mathbf{x}_t=\mathbf{x};\boldsymbol{\pi},\mathbf{P}\right) \tag{11}$$

where d^*_{t+j} represents the optimal employment decision j periods ahead under the assumption that the probabilities in $\mathbf{P}$ are the ones associated with the optimal decision rule. It is important to emphasize that to obtain the values $Z_{\boldsymbol{\pi},\mathbf{P}}(d,\mathbf{x})$ and $e_{\boldsymbol{\pi},\mathbf{P}}(d,\mathbf{x})$ we only need to know the probabilities $\mathbf{P}$, the parameters $\boldsymbol{\pi}$, and the discount factor β.

We provide here a brief sketch of how to calculate the values $Z_{\boldsymbol{\pi},\mathbf{P}}(d,\mathbf{x})$ and $e_{\boldsymbol{\pi},\mathbf{P}}(d,\mathbf{x})$. For more detailed, see Aguirregabiria and Mira (2002, 2009). Let $W_z^{\mathbf{P}}(\mathbf{x})\boldsymbol{\theta} + W_e^{\mathbf{P}}(\mathbf{x})$ be the expected discounted utility of behaving according to choice probabilities $\mathbf{P}$ from current period t and into the infinite future when $\mathbf{x}_t=\mathbf{x}$. Consider that the space of state variables X is discrete. Define the matrix $\mathbf{W}_z^{\mathbf{P}} \equiv \{W_z^{\mathbf{P}}(\mathbf{x}) : \mathbf{x}\in X\}$ and the vector $\mathbf{W}_e^{\mathbf{P}} \equiv \{W_e^{\mathbf{P}}(\mathbf{x}) : x\in X\}$. One can show that $\mathbf{W}_z^{\mathbf{P}}$ is the unique solution of the recursive equation $\mathbf{W}_z^{\mathbf{P}} = \sum_{d\in D}\mathbf{P}(d) * \{\mathbf{z}(d)+\beta\;\mathbf{F}_x(d)\;\mathbf{W}_z^{\mathbf{P}}\}$ where: $\mathbf{P}(d)$ is the column vector of choice probabilities $\{P(d|\mathbf{x}) : \mathbf{x}\in X\}$; $\mathbf{z}(d)$ is the matrix $\{z(d,\mathbf{x}) : \mathbf{x}\in X\}$; $\mathbf{F}_x(d)$ is a transition probability matrix with elements $f_x(\mathbf{x}_{t+1}|\mathbf{x}_t,d)$; and $*$ is the element-by-element product. Similarly, $\mathbf{W}_e^{\mathbf{P}}$ is the unique solution of the recursive equation $\mathbf{W}_e^{\mathbf{P}} = \sum_{d\in D}\mathbf{P}(d) * \{\mathbf{e}^{\mathbf{P}}(d)+\beta\;\mathbf{F}_x(d)\;\mathbf{W}_z^{\mathbf{P}}\}$ where $\mathbf{e}^{\mathbf{P}}(d)$ is the vector $\{e^{\mathbf{P}}(d,\mathbf{x}) : \mathbf{x}\in X\}$ and $e^{\mathbf{P}}(d,\mathbf{x}) \equiv E(\xi(d^*_t,\boldsymbol{\varepsilon}_t)|d^*_t=d,\mathbf{x}_t=\mathbf{x})$, that is a known function of the choice probabilities $P(d|\mathbf{x})$. One can compute $\mathbf{W}_z^{\mathbf{P}}$ and $\mathbf{W}_e^{\mathbf{P}}$ by successive approximations, iterating on the contraction mappings that implicitly define these objects.[9]

[9] Note that the mappings that implicitly define $\mathbf{W}_z^{\mathbf{P}}$ and $\mathbf{W}_e^{\mathbf{P}}$ are linear in these objects. Therefore, there is

Aguirregabiria and Mira (2002) derive the properties of the contraction mapping $\Psi_{\theta,\sigma_\varepsilon,\boldsymbol{\pi}}$. For a given vector of structural parameters $(\boldsymbol{\theta}, \sigma_\varepsilon, \boldsymbol{\pi})$ we can use this mapping to obtain the vector of optimal choice probabilities $\mathbf{P}$. This mapping is also an important component in the maximum likelihood estimation of $(\boldsymbol{\theta}, \sigma_\varepsilon)$ using the nested procedure proposed by Aguirregabiria and Mira (2002). We describe this procedure in section 5.3.

4.3 Assumptions for Policy Evaluation

Our main interest is to evaluate the effects of the 1984 labor market reform on employment and productivity in the Spanish manufacturing industry, using our longitudinal sample of Spanish firms. The reform extended the use of temporary contracts to any activity, temporary or not, and reduced firing costs for these contracts from 45 days to 12 days of wages per year of tenure. The new regulation applied to every type of firms and workers, regardless of size, industry, occupation, etc. Our approach for evaluating the effects of this reform exploits sample information before and after the policy change, together with the structure of our model and the characteristics of the reform. This section describes our evaluation approach and its identification assumptions.

Some identification assumptions are necessary to establish that the pre-reform and the post-reform sample periods correspond to the equilibria with the old and the new policies, respectively. This subsection discusses these assumptions.

a) Non anticipation of the reform. If agents would have anticipated the policy change, their behavior before the reform would not represent their optimal decisions if the reform would not have taken place. For instance, some firms willing to hire in 1983 or 1984 might have preferred to postpone hiring and firing decisions until 1985 in order to use the new type of labor contract. Departures from this assumption might bias our estimates of labor

a closed form expression for $\mathbf{W}_z^{\mathbf{P}}$ and for $\mathbf{W}_e^{\mathbf{P}}$. For instance, $\mathbf{W}_z^{\mathbf{P}} = (I - \beta \sum_{a=0}^{J} \mathbf{P}(a) * \mathbf{F}_x(a))^{-1} \sum_{a=0}^{J} \mathbf{P}(a) * \mathbf{z}(a)$. When the number of cells in X is small enough, matrix inversion algorithms may be preferable to successive approximations. The matrix $(I - \beta F)^{-1}$ can also be approximated using the series $I + \beta F + \beta^2 F^2 + ... + \beta^K F^K$, with K large enough. This can be easier than matrix inversion. More generally, this inverse matrix can be obtained iterating in A (succesive approximations) in the mapping $A = I + \beta F\ A$.

adjustment costs for the pre-reform period. Since our sample covers only three years before the policy change, there is not very much we can do to control for this potential source of bias.

b) Instantaneous learning about the features of the new policy. Looking at our data, there is clear evidence that a long transition period to the new steady-state took place after the reform. In particular, the proportion of temporary employment increased almost every year between 1984 and 1993, and was kept at high levels since then (see Figure 2). Such transition period could be explained by the existence of large firing costs for permanent workers. However, another reason that might have contributed to this long transition, and which is not considered in our model, is that firms could have learnt slowly about the features of the new policy rule. Since our model assumes instantaneous learning, it rules out this alternative explanation. The fact that the number of new temporary contracts exerted a large increase shortly after the reform, in 1985, seems to support our assumption.[10]

c) Policy-related and policy-invariant parameters. The reform entailed a change in the dismissal costs of temporary workers, θ_F^T. Hiring costs of temporary workers, θ_H^T, their relative productivity, λ, and hiring costs of permanent workers may have been affected by the reform as well. Therefore, θ_F^T, θ_H^T, θ_H^P and λ are policy related parameters. It seems plausible that the technological parameter α_L, firing costs of permanent workers, θ_F^P, and the stochastic process of the productivity shock are policy invariant parameters. In an equilibrium framework, the stochastic process of wages may be affected by this reform. This is consistent with the times series of wages for permanent and temporary workers in figure 7. Our estimation of the econometric model takes into account this possibility. However, our model is of partial equilibrium and our policy evaluation (i.e., counterfactual experiments) provides partial equilibrium effects.

[10] It is also important to emphasize that our assumption of instantaneous learning does not imply that the new steady-state was reached instantaneously after the reform. Our econometric approach assumes that the structural equations (i.e., the production function, the wage equation, and the dynamic labor demand functions) are stable within the pre-reform period and within the post-reform period. This assumption is fully consistent with the evidence on long transition periods for some endogenous variables.

5 Estimation of the model

We have a panel dataset with firm-level, annual-frequency information on output, employment by type of contract, physical capital, investment, and wage bill: $\{Y_{it}, L_{it}^T, L_{it}^P, K_{it}, I_{it}, WB_{it} : i = 1, ..., N;\ t = 1, ..., T_i\}$. Our econometric model consists on the production function, the stochastic processes for the productivity shock and wages, and the dynamic model for the demand of permanent and temporary labor. The vector of structural parameters is $\{\alpha_L, \lambda, \boldsymbol{\theta}, \sigma_\varepsilon, \boldsymbol{\pi}, \beta\}$, where $\boldsymbol{\theta}$ is (as defined in previous section) the vector of adjustment costs parameters, and $\boldsymbol{\pi}$ is the vector of parameters in the transition probabilities of the state variables. We estimate these parameters in two steps. In a first step, we estimate $\{\alpha_L, \lambda, \boldsymbol{\pi}\}$ from the production function and transition data. In a second step, we estimate $\boldsymbol{\theta}$ and σ_ε by maximum likelihood in the dynamic labor demand model. Before we describe our estimation methods and results, we explain first our approach to estimate wages by type of contract.

5.1 Estimation of wages

As we mentioned in Section 3, our sample information on the firm's total wage bill is not broken down by type of contract. By definition, the wage bill of firm i at year t is $WB_{it} = W_{it}^T L_{it}^T + W_{it}^P L_{it}^P$. Given our assumption that the wage of temporary workers is the same for every firm, we have that:

$$\frac{WB_{it}}{L_{it}^P} = W_t^T \left(\frac{L_{it}^T}{L_{it}^P}\right) + W_{it}^P \tag{12}$$

We observe WB_{it}/L_{it}^P and L_{it}^T/L_{it}^P. But we do not have data on W_t^T and W_{it}^P, at least for every year in our sample period and at the firm-level. Therefore, W_t^T and W_{it}^P are unobservables for us. If the wage of permanent workers were mean independent of the temporary-to-permanent ratio, L_{it}^T/L_{it}^P, we could estimate the value W_t^T by running a regression for WB_{it}/L_{it}^P on L_{it}^T/L_{it}^P (interacted with time dummies). Moreover, the residual of that regression would be a consistent estimator of the wage of permanent workers at the firm level. However, such estimate of W_t^T will be affected by an upward endogeneity bias if, as we expect, the temporary-to-permanent ratio is positively correlated with the wage of permanent workers.

To control for this bias, we consider a fixed-effect or within-firms estimator. That is, we assume that the wage of permanent workers is:

$$W_{it}^{P} = \mu_i + \gamma_t + u_{it} \tag{13}$$

where μ_i is a firm fixed-effect; γ_t is an aggregate effect; and u_{it} is a shock assumed to be uncorrelated with the temporary-to-permanent ratio. Under this assumption, the fixed-effects estimator provides consistent estimates of W_t^T.

Figure 7 presents the time series of our fixed effect estimates for the average wages of permanent and temporary workers. According to our estimates, the wage differential between contracts was small before the reform but it has widened very importantly after 1984. This result is consistent with the evidence provided by Bentolila and Dolado (1994). As argued by these authors, a possible explanation for this wage differential is that the own existence of temporary contracts increased the job security and the wage bargaining power of permanent workers.

5.2 Estimation of the production function

The specification of the production function in equation (2) treats physical capital as a component of the productivity shock η_{it}. This is a convenient assumption to reduce the dimensionality of decision and state spaces. Though we maintain this assumption throughout the paper, in the estimation of the production function we incorporate explicitly physical capital and estimate the technological parameter associated with this input. Looking at this estimate is a way of checking for the validity of the specification and for the economic sense of the estimation results. We consider a Cobb-Douglas production function in terms of physical capital and production-equivalent units of labor. The production function in logarithms is:

$$\ln Y_{it} = \alpha_K \ \ln K_{it} + \alpha_L \ \ln(L_{it}^{P} + \lambda \ L_{it}^{T}) + \omega_{it} \tag{14}$$

where K_{it} is the installed capital at the beginning of year t; ω_{it} is the "pure" productivity shock such that $\eta_{it} = \alpha_K \ \ln K_{it} + \omega_{it}$.

It is well known that the OLS estimation of this equation may suffer of endogeneity bias because correlation between the values of inputs and the unobservable productivity shock. Furthermore, if the productivity shock is serially correlated, lagged values of inputs and output are also correlated with the unobservables, and therefore they cannot be used as instruments. Using input prices (e.g., wages) as instruments is also problematic. Some input prices do not have variability at the firm level (e.g., the wage of temporary workers, or the price of capital), and those prices that do have that variability are very suspicious of being correlated with firm's productivity (e.g., the wage of permanent workers).

Our identification of the parameters in the production function is based on the control function approach proposed by Olley and Pakes (1996). Our application of this method is in the spirit of the extension proposed by Ackerberg, Caves and Frazer (2006). Investment in physical capital is a function of the state variables $(K_{it}, L^P_{i,t-1}, L^T_{i,t-1}, C_t, \omega_{it})$, where C_t represents input prices, and of some shocks χ_{it} which are assumed to be independent of the other state variables $(K_{it}, L^P_{i,t-1}, L^T_{i,t-1}, C_t, \omega_{it}, \varepsilon_{it})$. Let $I_{it} = g(K_{it}, L^P_{i,t-1}, L^T_{i,t-1}, C_t, \omega_{it}, \chi_{it})$ be the optimal decision rule for investment. Since this function g is strictly increasing in the productivity shock ω_{it}, there is an inverse function such that $\omega_{it} = g^{-1}(I_{it}, K_{it}, L^P_{i,t-1}, L^T_{i,t-1}, C_t, \chi_{it})$. Based on this expression, we can decompose ω_{it} in two additive terms: $\omega_{it} = \omega^e_{it} + \chi^*_{it}$, where $\omega^e_{it} \equiv E\left(\omega_{it} | I_{it}, K_{it}, L^P_{i,t-1}, L^T_{i,t-1}, C_t\right)$ and χ^*_{it} is the remaining part of ω_{it}. This decomposition has two important features. First, ω^e_{it} only depends on observable variables. And second, χ^*_{it} is, by construction, mean independent of $\left(I_{it}, K_{it}, L^P_{i,t-1}, L^T_{i,t-1}, C_t\right)$, and also of L^P_{it} and L^T_{it}. Therefore, we can write the production function as,

$$\ln Y_{it} = \alpha_L \ \ln(L^P_{it} + \lambda \ L^T_{it}) + \eta^e_{it} + \chi^*_{it} \tag{15}$$

where $\eta^e_{it} = \alpha_K \ln K_{it} + \omega^e_{it}$. Note that η^e_{it} is a smooth function of $(I_{it}, K_{it}, L^P_{i,t-1}, L^T_{i,t-1}, C_t)$. We can control for this term by including a high order polynomial in these observable variables. The key identification assumption is that there are i.i.d. shocks ε_{it} and χ_{it} affecting current employment and investment, respectively, which are mutually independent. Under this assumption, we can use current investment to control for the endogenous part of the

productivity shock ω_{it}, and still we have some variability left in the current employment variables L_{it}^P and L_{it}^T to identify α_L and λ.

Once we have estimated α_L and λ, we can exploit the assumption on the Markov process of ω_{it} to estimate α_K. First, we obtain estimates of η_{it} as the residuals $\ln Y_{it} - \widehat{\alpha}_L \ln(L_{it}^P + \widehat{\lambda} L_{it}^T)$. According to the model, $\eta_{it} = \alpha_K \ln K_{it} + \omega_{it}$. Assuming that ω_{it} follows an AR(1) process: $\omega_{it} = \rho_\omega \, \omega_{i,t-1} + a_{it}$ with $a_{it} \sim iid\,(0, \sigma_a^2)$, we have that

$$\left(\eta_{it} - \rho_\omega \, \eta_{i,t-1}\right) = \alpha_K \, \left(\ln K_{it} - \rho_\omega \, \ln K_{i,t-1}\right) + a_{it} \tag{16}$$

The innovation a_{it} is independent of $\eta_{i,t-1}$, $\ln K_{it}$ and $\ln K_{i,t-1}$, and therefore we can estimate α_K and ρ_ω using least squares, or the Cochrane-Orcutt iterative procedure.

In Table 3, we present our estimates of the production function parameters. For the sake of comparison, we report estimates using both the Olley-Pakes method and the (inconsistent) nonlinear least squares estimator. All the estimations include time dummies and 20 industry dummies. The control function η_{it}^e includes all the terms of a second order polynomial in $(I_{it}, K_{it}, L_{i,t-1}^P, L_{i,t-1}^T)$ and interactions of these terms with time dummies, what entails a total of 164 regressors. The parameters λ, ρ_ω and σ_a are allowed to change between the pre-reform and the post-reform period. However, whereas a change in λ might be attributed to the reform, changes in ρ_ω or in σ_a might not. Comparing the two reported estimates, both the magnitudes and the qualitative results are fairly similar, the major differences concerning the λ parameter before the reform.

The point estimates imply some decreasing returns to scale, though the hypothesis of constant returns to scale cannot be rejected under typical significance levels. The most interesting result in this table is the post-reform increase in the relative efficiency of temporary labor. While this input was just half as efficient as permanent labor before 1984, it has become almost as efficient after the reform. A possible explanation for this result is that adverse selection was a more serious problem for temporary labor in the pre-reform period. However, we should be cautious to attribute this parameter change entirely to the reform. For instance, young workers in Spain during this period were significantly more ed-

ucated than older cohorts, and they have also accounted for a large proportion of temporary contracts. The estimates of the parameters ρ_ω and σ_a before and after the reform suggest small reductions in the persistence of the productivity shock and in the variability of the innovation.

5.3 Estimation of the dynamic labor demand model

We estimate the dynamic labor demand model using the Nested Pseudo Likelihood (NPL) algorithm proposed by Aguirregabiria and Mira (2002). The NPL is a procedure to estimate discrete choice dynamic programming models that, in the context of single-agent models, provides the maximum likelihood estimator of the structural parameters. We provide here a description of this procedure in the context of our model. In this section, we treat the variables W_t^T, W_{it}^P and η_{it} as observable to the researcher. These variables in fact has been consistently estimated in a first step, and therefore we really observe the estimated values $\hat{W}_t^T$, $\hat{W}_{it}^P$ and $\hat{\eta}_{it}$. For notational convenience, we omit the 'hats'. The fact that the estimated values include estimation error does not affect the consistency of our estimator of $\boldsymbol{\theta}$, though it affects its asymptotic variance.

Let $P_0(d_{it}|\mathbf{x}_{it})$ be the true distribution of employment changes, $d_{it} \equiv \{\Delta L_{it}^P, \Delta L_{it}^T\}$, conditional on the state variables, $\mathbf{x}_{it} \equiv (L_{it-1}^P, L_{it-1}^T, W_t, \eta_{it})$, in the population of our study. Define the vector $\mathbf{P}_0 \equiv \{P_0(d|\mathbf{x}) : (d, \mathbf{x}) \in D \times X\}$. And define the (pseudo) log-likelihood function:

$$Q\left(\boldsymbol{\theta}, \sigma_\varepsilon, \boldsymbol{\pi}, \mathbf{P}_0\right) \;=\; \textstyle\sum_{i=1}^N \sum_{t=1}^{T_i} \ln \Psi_{\boldsymbol{\theta},\sigma_\varepsilon,\boldsymbol{\pi}}(\mathbf{P}_0)(d_{it}|\mathbf{x}_{it}) \tag{17}$$

where

$$\Psi_{\theta,\sigma_\varepsilon,\boldsymbol{\pi}}(\mathbf{P})(d|\mathbf{x}) =$$

$$\int \frac{\exp\left\{Z_{\boldsymbol{\pi},\mathbf{P}}\left(d,\mathbf{x}\right)\frac{\theta}{\sigma_0} + e_{\boldsymbol{\pi},\mathbf{P}}\left(d,\mathbf{x}\right) + \varepsilon^P \Delta L^P(d)\frac{\sigma_P}{\sigma_0} + \varepsilon^T \Delta L^T(d)\frac{\sigma_T}{\sigma_0}\right\}}{\sum\limits_{j\in D} \exp\left\{Z_{\boldsymbol{\pi},\mathbf{P}}\left(j,\mathbf{x}\right)\frac{\theta}{\sigma_0} + e_{\boldsymbol{\pi},\mathbf{P}}\left(j,\mathbf{x}\right) + \varepsilon^P \Delta L^P(j)\frac{\sigma_P}{\sigma_0} + \varepsilon^T \Delta L^T(j)\frac{\sigma_T}{\sigma_0}\right\}} \phi(d\varepsilon^P)\phi(d\varepsilon^T) \tag{18}$$

Let $\hat{\mathbf{P}}_0$ be a nonparametric estimator of the set of conditional choice probabilities in $\mathbf{P}_0$. And let $\hat{\boldsymbol{\pi}}$ be an estimator of the parameters in the transition probability functions of wages and the productivity shock. Given these estimates, we can calculate the values $Z_{\hat{\boldsymbol{\pi}},\hat{\mathbf{P}}_0}(d, \mathbf{x}_{it})$ and $e_{\hat{\boldsymbol{\pi}},\hat{\mathbf{P}}_0}(d, \mathbf{x}_{it})$ using the recursive method that we described in section 4.2. Therefore, given $Z_{\hat{\boldsymbol{\pi}},\hat{\mathbf{P}}_0}(d, \mathbf{x}_{it})$ and $e_{\hat{\boldsymbol{\pi}},\hat{\mathbf{P}}_0}(d, \mathbf{x}_{it})$, the function $Q(\boldsymbol{\theta}, \sigma_\varepsilon, \hat{\boldsymbol{\pi}}, \hat{\mathbf{P}}_0)$ is the log-likelihood of a random-coefficients multinomial logit model, where the random coefficients come from the term $\varepsilon^P \Delta L^P(d)\frac{\sigma_P}{\sigma_0} + \varepsilon^T \Delta L^T(d)\frac{\sigma_T}{\sigma_0}$.[11] Given this likelihood, we can estimate the parameters $\boldsymbol{\theta}$, σ_0, σ_P, and σ_T. Note that these parameters are separately identified from $\boldsymbol{\theta}/\sigma_0$, σ_T/σ_0, and σ_P/σ_0 because the first element of $\boldsymbol{\theta}$, which is associated with the value of output minus the wage bill, is equal to 1. The estimator of $(\boldsymbol{\theta}, \sigma_\varepsilon)$ that maximizes $Q(\theta, \sigma_\varepsilon, \hat{\boldsymbol{\pi}}, \hat{\mathbf{P}}_0)$ is consistent and asymptotically equivalent to the maximum likelihood estimator (see Proposition 4 in Aguirregabiria and Mira, 2002). Furthermore, as shown by Aguirregabiria and Mira (2002), asymptotic standard errors do not have to be corrected for the estimation error in $\hat{\mathbf{P}}_0$. A recursive extension of this two-step method returns the (conditional) maximum likelihood estimator of $(\boldsymbol{\theta}, \sigma_\varepsilon)$. The main computational and econometric issues in this estimation procedure concerns the computation of the values $Z_{\hat{\boldsymbol{\pi}},\hat{\mathbf{P}}_0}(d, \mathbf{x}_{it})$ and $e_{\hat{\boldsymbol{\pi}},\hat{\mathbf{P}}_0}(d, \mathbf{x}_{it})$.

(a) Discretization of employment changes (decision variable) and of employment levels (endogenous state variables). As mentioned above, the main reason why we consider a discrete model is that there is significant lumpiness in employment changes. For more than 57% of the firm-year observations in the sample, the annual change in temporary and permanent employment is between -5 and $+5$ workers, and more than 72% of the observations lie between -10 and $+10$ workers. Table 4 presents the empirical distribution of employment changes before and after the reform. We can see that a small number of discrete values account for a large proportion of observations of employment changes. However, though the distribution of employment changes is discrete and lumpy, it also has long tails. We would

[11] We calculate numerically the double integral in the probability function $\Psi_{\theta,\sigma_\varepsilon,\pi}(P)(d|x)$. More specifically, we use the Gauss-Legendre quadrature method provided by the command *intquad2* in the GAUSS software package.

need a support with too many values to account for more than 90% of the sample values of this variable. Similarly, there is a trade-off in the discretization of the endogenous state variables L^P_{it-1} and L^T_{it-1}. A finer discretization can capture more sample variation of the variables, but it also increases the cost of computing the present values $Z_{\hat{\boldsymbol{\pi}},\hat{\mathbf{P}}_0}$ and $e_{\hat{\boldsymbol{\pi}},\hat{\mathbf{P}}_0}$.[12] Also, the discretizations of ΔL^j_{it} and L^j_{it-1} should be consistent with each other.

Taking into account these issues, we consider the following approach. Define the variable $\ell^j_{it} \equiv 100 * (L^j_{it}/\bar{L}_i)$, where $\bar{L}_i$ is the sample mean of total employment $L^P_{it} + L^T_{it}$ for firm i. Therefore, ℓ^j_{it} represents the percentage of current employment (type j) relative to the firm specific mean. Define also $d^j_{it} \equiv 100 * (\Delta L^j_{it}/\bar{L}_i)$, that measures the percentage of current employment change (type j) relative to the firm-specific mean. It is clear that:

$$\ell^j_{it} = \ell^j_{i,t-1} + d^j_{it} \tag{19}$$

We discretize the space of d^P_{it} (and d^T_{it}) in the set of integer numbers multiples of 2 between -20 and $+20$. Note that $d^P_{it} = 0$ and $d^T_{it} = 0$ represents actual zeros in employment change.[13] The discretized space for ℓ^P_{it} (ℓ^T_{it}) is the set of integer numbers multiples of 2 between 40 and 120 (between 0 and 40). Figure 8 presents the histograms of the discretized values of the decision variables d^P_{it} and d^T_{it}, and of the state variables ℓ^P_{it} and ℓ^T_{it}.

(b) Discretization of exogenous state variables. We follow Tauchen (1986) and Tauchen and Hussey (1991) for the choice of the discretization grid of the exogenous state variables (W^T_t,W^P_{it}, η_{it}). For each of these variables, we estimate an AR(1) process and follow Tauchen-Hussey procedure. However, for the state variables (W^P_{it}, η_{it}) we apply a different discretization for each individual firm. That is, the discretization applies to the variables in deviations with respect to their firm-specific means: $W_{it} - \bar{W}_i$ and $\eta_{it} - \bar{\eta}_i$. By using firm-specific discretizations, we can capture most of the time-series variability of the state variables without having to consider too many grid points. The total number of cells in

[12] As explained in section 4.2, to obtain these present values we have to solve for $\mathbf{W}^{\mathbf{P}}_z$ in the system of equations $\mathbf{W}^{\mathbf{P}}_z = \sum_{d\in D} \mathbf{P}(d) * \{\mathbf{z}(d) + \beta\ \mathbf{F}_x(d)\ \mathbf{W}^{\mathbf{P}}_z\}$. The dimension of this system of equations is the number of cells in the state space X.

[13] Values greater (lower) than zero but lower than 2 (greater than -2) are censored at 2 (-2).

the discretized state space X is $6,888$ (i.e., 41 for permanent employment, 21 for temporary employment, 2 for wage of temporaries, 2 for wage of permanents, and 2 for the productivity shock).[14]

(c) Initial estimates of conditional choice probabilities. We have estimated a multinomial logit with dependent variable $d_{it} = (d_{it}^P, d_{it}^T)$, using as explanatory variables the terms of a second order polynomial in the state variables $(\ell_{it}^P, \ell_{it}^T, W_t^T, W_{it}^P, \eta_{it})$.

We have estimated our dynamic labor demand model under different specifications of the unobservables, including the pure conditional logit without random coefficients, and different random-coefficient models with homoscedastic and with heteroscedastic $\varepsilon's$. The time-discount factor is fixed at $\beta = 0.95$. For the selection of our most-preferred specification, we have considered two criteria: the sign and magnitude of the estimated parameters should have economic sense; and the model should provide a reasonable fit for aggregate statistics such as the aggregate time path of the proportion of temporary workers, the percentage of zeroes in the distribution of employment changes, average job turnover rates, and the cross-sectional variance of employment levels. Following these criteria, our favorite specification is a model where labor adjustment costs (fixed, linear and quadratic) and the standard deviation of the unobservable $\varepsilon's$ are proportional to the firm-specific mean of the salary-per-worker. For instance, the linear cost of firing permanent workers for firm i is $\theta_{F1i}^P = \tilde{\theta}_{F1}^P \ \bar{W}_i$, where $\tilde{\theta}_{F1}^P$ is the same parameter for every firm and $\bar{W}_i$ is firm $i's$ mean salary per worker, i.e., $\bar{W}_i = (1/T_i) \sum_{t=1}^{T_i} W_{it}$. The same specification applies to the other θ parameters in labor adjustment costs. Similarly, the variances of the unobservables are $var(\varepsilon_{it}^P) = \sigma_P^2 \ \bar{W}_i^2$, $var(\varepsilon_{it}^T) = \sigma_T^2 \ \bar{W}_i^2$, and $var(\varepsilon_{it}^0) = \sigma_0^2 \ \bar{W}_i^2$. It is important to note that

[14] In our computation of the inclusive values $Z_{\hat{\boldsymbol{\pi}},\hat{\mathbf{P}}_0}$ and $e_{\hat{\boldsymbol{\pi}},\hat{\mathbf{P}}_0}$, there is a technical issue and a simplifying assumption that deserve to be explained. First, given our definition of the state variables in deviations with respect to firm-specific means, the matrix $\mathbf{W}_z^{\mathbf{P}}$ and the vector $\mathbf{W}_e^{\mathbf{P}}$ are firm-specific, and they should be calculated on a firm-by-firm basis. This means that we have to solve $2,356$ systems of linear equations, each with a dimension of $6,888$ variables. Note that, for the implementation of our procedure, we do not need to store in memory the $2,356$ matrices $\mathbf{W}_z^{\mathbf{P}}$ and vectors $\mathbf{W}_e^{\mathbf{P}}$. We only need to store in memory one at a time. Second, for the calculation of $e_{\hat{\boldsymbol{\pi}},\hat{\mathbf{P}}_0}$ we have used the simplifying assumption that the future values of ε^P and ε^T are equal to the expected values. Therefore, the only component in $e_{\hat{\boldsymbol{\pi}},\hat{\mathbf{P}}_0}$ is the one that comes from the expectation of the future extreme value error ε^0.

the model with random coefficients provides both more sensible results and better fit that the pure conditional logit model. For instance, under the conditional logit model, the estimates of some lump-sum adjustment costs are negative and significant, and most quadratic adjustment costs are unrealistically large. That model cannot fit the thick tails in distribution of employment changes that we observe in the data.

Table 5 presents the estimates of the dynamic labor demand model for our preferred specification. We have estimated the model for three sub-periods: the pre-reform period 1983-1984, and the post-reform periods 1985-1988, and 1989-1992. Table 6 provides measures of goodness of fit of the estimated model.

Remark 1. Table 6 shows that the model provides a very good fit of different statistics that represent the cross-sectional distribution and the dynamics of permanent and temporary employment. The proportion of zeroes in the change of temporary employment is not fitted as well as the other statistics, i.e., the model under-estimates this proportion. However, the under-estimation of this statistic is similar for pre-reform and post-reform periods.

Remark 2. The most important part of labor adjustment costs is the linear component. That is the case for both permanent and temporary employment, and for hiring and firing. Quadratic costs are very small and not significant. Most fixed costs are also small, with the exception of the cost of firing of permanent workers and, very particularly, the cost hiring temporaries in the pre-reform period.

Remark 3. Linear (per worker) hiring costs are similar for temporary and for permanent workers. They are between 10% and 18% of a worker's annual salary. It seems that the per-worker hiring cost for permanent employment has declined after the reform. A possible interpretation of this result is that, after the reform, hiring a permanent employee typically means promoting a temporary worker to a permanent position. The promotion of an insider may be less expensive than the recruitment of an outsider.

Remark 4. Linear (per worker) firing costs of permanent workers are sizeable. They are between 46% and 53% of a worker's annual salary. They have been quite stable between

the pre-reform and the post-reform periods. Linear firing costs of temporary workers are relatively small (between 4% and 10%) and statistically not significant. They have declined between the pre-reform and the post-reform periods.

Remark 5. The most significant change between the pre-reform and the post-reform period is the very large reduction in the fixed cost of hiring temporary workers. Other significant changes are a drop in the linear costs of hiring permanent and temporary workers, and a decline in the linear cost of firing.

Remark 6. The parameters that measure the dispersion of the unobservable shocks are all positive, significant, and quite stable over time.

From our estimation of the structural equations, the overall picture that appears on the effects of the reform is the following: (1) it has made it cheaper to hire and fire temporary workers, both at the intensive and at the extensive margin; (2) it has reduced the cost of hiring permanent workers, probably because promoting an insider (temporary) to a permanent position is less expensive than recruiting an outsider; and (3) the productivity of temporary workers has become closer to the one of permanents; and (4) the wage-gap between permanent and temporary workers is now wider than before the reform.

6 Policy evaluation

We use the estimated model to evaluate the effects, on employment, job turnover, productivity and firms' value, of the introduction of temporary contracts. We also compare these effects with those associated with a counterfactual policy that halves linear-firing costs for all type of workers. To implement these policy evaluations, we select the firms active in the sample in year 1984. For this group of firms, we solve for the value function and the optimal decision rule in three dynamic programming models: a "pre-reform" model, a "post-reform" model, and a counterfactual model. For the three models, wages are assumed to be constant at their 1984 levels (i.e., the policy evaluation considers partial equilibrium effects), and the stochastic process of the productivity shock is the one for the period 1983-1984. In the pre-

reform model, the value of the other structural parameters are the ones estimated for the period 1983-1984. For the post-reform model, the structural parameters are the estimates for the period 1989-1992. Finally, for the counterfactual model, we fixed the values of the parameters at their 1983-1984 level, except for the linear firing costs θ_{F1}^{P} and θ_{F1}^{T} which are reduced by half: i.e., the counterfactual values of $\theta_{F1,i}^{P}/\bar{W}_i$ and $\theta_{F1,i}^{T}/\bar{W}_i$ are 0.257 and 0.049, respectively. For each model, we calculate the steady-state distribution of the state variables and use this distribution to obtain the mean values of employment, output, etc.

Table 7 presents the results of these experiments. The introduction of temporary contracts had important positive effects on total employment (a 3.5% increase) and job turnover. The increase in total employment is associated with a strong substitution of permanent by temporary workers: the proportion of temporary workers goes from 3.8% to 16.2%. Permanent employment declines by 10%. The positive effects on productivity (0.7%) and the value of firms (1.2%) are small. These effects contrast substantially with the ones of the counterfactual reform. While the effects on total employment are similar (a 4.1% increase), the counterfactual reform improves permanent employment (6.6% increase), labor productivity (1.9% increase), and the value of firms (4.8%). Furthermore, the proportion of temporary employment becomes almost null (1.3%).

7 Concluding remarks

Using panel data of Spanish manufacturing firms, we have estimated a dynamic labor demand model and evaluated the effects of a reform that introduced temporary contracts in 1984. The structural model allows for a rich specification of labor adjustment costs, including fixed, linear and quadratic components, and unobserved firm-heterogeneity (i.e., random coefficients). The model with random-coefficients provides a better fit and more sensible results that a simpler conditional logit model. Our estimation results show significant changes in structural parameters after the reform. Hiring and firing temporary workers has become less expensive, both at the intensive and at the extensive margins, and the cost of hiring

permanent workers has declined. Based on the estimated model, we present counterfactual experiments to evaluate the effects of the reform. We find important effects on employment and job turnover, but modest effects on productivity and value of firms. However, we also find that a counterfactual policy that halves firing costs for all contracts has similar effects on total employment, but significantly stronger positive effects on output, value of firms, and permanent employment.

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Table 1
Distribution of firms by 2-digit industry and by size
Unbalanced panel 1982-1993 (2356 firms)

		Small	*Med1*	*Med 2*	*Large*	*Total*
Iron, steel	Abs. freq.	5	8	10	22	45
and metal	% by ind.	11.11	17.78	22.22	48.89	100.00
(22)	% by size	1.29	0.94	1.73	4.10	1.91
Bldg. materials	Abs. freq.	27	88	34	33	182
glass, ceramics	% by ind.	14.84	48.35	18.68	18.13	100.00
(24)	% by size	6.98	10.29	5.89	6.15	7.72
Chemicals	Abs. freq.	39	99	76	92	306
	% by ind.	12.75	32.35	24.84	32.07	100.00
(25)	% by size	10.08	11.58	13.17	17.13	12.99
Non-ferrous	Abs. freq.	38	103	53	31	225
metal	% by ind.	16.89	45.78	23.56	13.78	100.00
(31)	% by size	9.82	12.05	9.19	5.77	9.55
Basic	Abs. freq.	29	52	47	33	161
machinery	% by ind.	18.01	32.30	29.19	20.50	100.00
(32)	% by size	7.49	6.08	8.15	6.15	6.83
Office	Abs. freq.	0	1	0	3	4
machinery	% by ind.	0.00	25.00	0.00	75.00	100.00
(33)	% by size	0.00	0.12	0.00	0.56	0.17
Electric	Abs. freq.	11	29	24	35	99
materials	% by ind.	11.11	29.29	24.24	35.35	100.00
(34)	% by size	2.84	3.39	4.16	6.52	4.20
Electronic	Abs. freq.	3	8	10	14	35
	% by ind.	8.57	22.86	28.57	40.00	100.00
(35)	% by size	0.78	0.94	1.73	2.61	1.49
Motor vehicles	Abs. freq.	8	21	25	36	13
	% by ind.	8.89	23.33	27.78	40.00	100.00
(36)	% by size	2.07	2.46	4.33	6.70	3.82
Ship	Abs. freq.	3	2	2	6	13
building	% by ind.	23.08	15.38	15.38	46.15	100.00
(37)	% by size	0.78	0.23	0.35	1.12	0.55
Other	Abs. freq.	2	5	5	6	18
motor vehicles	% by ind.	11.11	27.78	27.78	33.33	100.00
(38)	% by size	0.52	0.58	0.87	1.12	0.76
Precision	Abs. freq.	2	8	3	4	17
instruments	% by ind.	11.76	47.06	17.65	23.53	100.00
(39)	% by size	0.52	0.94	0.52	0.74	0.72

Table 1 (cont.)
Distribution of firms by 2-digit industry and by size
Unbalanced panel 1982-1993 (2356 firms)

		Small	*Med1*	*Med 2*	*Large*	*Total*
Non-elaborated	Abs. freq.	23	83	46	48	230
food	% by ind.	23.04	36.09	20.00	20.87	100.00
(41)	% by size	13.70	9.71	7.97	8.94	9.76
Food, tobacco	Abs. freq.	53	51	31	45	180
and drinks	% by ind.	29.44	28.33	17.22	25.00	100.00
(42)	% by size	13.70	5.96	5.37	8.38	7.64
Basic	Abs. freq.	20	57	53	37	167
Textile	% by ind.	11.98	34.13	31.74	22.16	100.00
(43)	% by size	5.17	6.67	9.19	6.89	7.09
Leather	Abs. freq.	4	16	12	4	36
	% by ind.	11.11	44.44	33.33	11.11	100.00
(44)	% by size	1.03	1.87	2.08	0.74	1.53
Garment	Abs. freq.	11	48	34	22	115
	% by ind.	9.57	41.74	29.57	19.13	100.00
(45)	% by size	2.84	5.61	5.89	4.10	4.88
Wood and	Abs. freq.	21	45	26	8	100
furniture	% by ind.	21.00	45.00	26.00	8.00	100.00
(46)	% by size	5.43	5.26	4.51	1.49	4.24
Cellulose and	Abs. freq.	29	63	42	33	167
paper edition	% by ind.	17.37	37.72	25.15	19.76	100.00
(47)	% by size	7.49	7.37	7.28	6.15	7.09
Plastic	Abs. freq.	22	46	33	17	118
materials	% by ind.	18.64	38.98	27.97	14.41	100.00
(48)	% by size	5.68	5.38	5.72	3.17	5.01
Other	Abs. freq.	7	22	11	8	48
non-basic	% by ind.	14.58	45.83	22.92	16.67	100.00
(49)	% by size	1.81	2.57	1.91	1.49	2.04
Total	Abs. freq.	387	855	577	537	2356
	% by ind.	16.43	36.29	24.49	22.79	100.00
	% by size	100.00	100.00	100.00	100.00	100.00

Note: Small means firm's time average of total employment lower or equal than 25. Med 1 means firm's time average of total employment greater than 25 and lower or equal than 75. Med 2 means firm's time average of total employment greater than 75 and lower or equal than 200. Large means firm's time average of total employment greater than 200.

Table 2
Descriptive Statistics Balanced panel 1982-1992 (389 firms)

Variable	**Period 1982-1984**			**Period 1989-1992**		
	Pctile 25	*Median*	*Pctile 75*	*Pctile 25*	*Median*	*Pctile 75*
Growth Real Output	-6.2%	2.4%	10.5%	-7.1%	2.3%	11.5%
Growth Total Employment	-3.6%	-0.6%	2.6%	-5.6%	-0.6%	4.4%
Number of Workers	60	131	297	65	137	298
Permanent Workers	55	128	276	56	121	272
Temporary Workers	0	0	3	0	6	22
% Temp Workers	0.0%	0.0%	2.2%	0.0%	4.9%	13.6%
Ratio (Sales / Wage Bill)	4.2	5.7	8.5	4.3	5.6	7.8
Number of observations		1167			1556	

Table 3
Estimation of Production Function Parameters
Unbalanced panel 1982-1993 (2356 firms)[1]

Parameters	**Least Squares** Estimate	(S.E.)[2]	**Olley-Pakes** Estimate	(S.E.)[2]
α_K	0.260	(0.006)	0.294	(0028)
α_L	0.690	(0.008)	0.680	(0.036)
Pre-Reform λ	0.666	(0.093)	0.549	(0.150)
Post-Reform λ	0.895	(0.035)	0.913	(0.054)
Pre-Reform ρ_ω	0.955	(0.010)	0.957	(0.011)
Post-Reform ρ_ω	0.931	(0.003)	0.943	(0.003)
Pre-Reform $\boldsymbol{\sigma}_a$	0.174	(-)	0.172	(-)
Post-Reform $\boldsymbol{\sigma}_a$	0.207	(-)	0.204	(-)
# Observations[3]	16,640		15,985	

Note (1): All the estimations include time dummies and 20 industry dummies.
Note (2): Standard errors are robust of heterocedasticity and autocorrelation.
Note (3): In Olley-Pakes estimation we can use only those observations with investment different than zero. This explains the smaller number of observations.

Table 4
Distribution of Employment Changes. Unbalanced panel

PRE-REFORM PERIOD: 1982-1984

		Change in Temporary Employment							
	%	**≤-3**	**-2**	**-1**	**0**	**+1**	**+2**	**≥+3**	**Total**
Change in Permanent Employment	**≤-3**	3.2	0.5	2.2	**22.8**	2.2	1.1	6.0	37.9
	-2	0.3	0.3	0.4	**3.8**	0.4	0.2	0.5	5.9
	-1	0.4	0.1	0.3	**6.5**	0.5	0.4	0.5	8.8
	0	**1.1**	**0.2**	**0.6**	**9.5**	**1.3**	**0.6**	**1.8**	**15.1**
	+1	0.6	0.1	0.5	**4.6**	0.8	0.4	0.6	7.4
	+2	0.5	0.1	0.2	**2.5**	0.3	0.2	0.8	4.4
	≥+3	2.3	0.5	0.7	**11.7**	0.9	0.4	3.8	20.4
	Total	8.5	1.7	4.9	**61.3**	6.3	3.4	13.9	100.0

POST-REFORM PERIOD: 1989-1992

		Change in Temporary Employment							
	%	**≤-3**	**-2**	**-1**	**0**	**+1**	**+2**	**≥+3**	**Total**
Change in Permanent Employment	**≤-3**	5.5	1.1	1.5	**10.1**	1.7	1.2	7.9	28.9
	-2	0.8	0.2	0.3	**2.6**	0.7	0.3	1.1	5.9
	-1	0.9	0.3	0.7	**4.1**	0.7	0.4	1.3	8.4
	0	**1.5**	**0.5**	**1.4**	**11.3**	**1.6**	**1.0**	**2.3**	**19.6**
	+1	1.0	0.3	0.8	**3.2**	0.9	0.4	1.2	7.7
	+2	0.5	0.4	0.5	**2.6**	0.5	0.3	0.9	5.8
	≥+3	6.3	0.7	0.8	**7.7**	1.0	1.1	6.0	23.7
	Total	16.5	3.5	6.0	**41.5**	7.0	4.7	20.7	100.0

Table 5
Estimation of of the Dynamic Labor Demand Model
Unbalanced panel 1982-1993 (2356 firms)[1]

Parameters		Period 1982-1984 Estimate (Std.Error)	Period 1985-1988 Estimate (Std.Error)	Period 1989-1992 Estimate (Std.Error)
Fixed Hiring Cost Perm:	$\frac{\boldsymbol{\theta}^{\mathbf{P}}_{\mathbf{H0,i}}}{\bar{W}_i}$	0.012 (0.061)	0.018 (0.035)	0.028 (0.041)
Linear Hiring Cost Perm:	$\frac{\boldsymbol{\theta}^{\mathbf{P}}_{\mathbf{H1,i}}}{\bar{W}_i}$	**0.183**** **(0.058)**	**0.117**** **(0.038)**	**0.101**** **(0.043)**
Quad Hiring Cost Perm:	$\frac{\boldsymbol{\theta}^{\mathbf{P}}_{\mathbf{H2,i}}}{\bar{W}_i}$	0.00031 (0.00060)	0.00063 (0.00078)	0.00055 (0.00052)
Fixed Firing Cost Perm:	$\frac{\boldsymbol{\theta}^{\mathbf{P}}_{\mathbf{F0,i}}}{\bar{W}_i}$	0.083** (0.038)	0.136** (0.024)	0.080** (0.036)
Linear Firing Cost Perm:	$\frac{\boldsymbol{\theta}^{\mathbf{P}}_{\mathbf{F1,i}}}{\bar{W}_i}$	**0.514**** **(0.098)**	**0.464**** **(0.035)**	**0.528**** **(0.080)**
Quad Firing Cost Perm:	$\frac{\boldsymbol{\theta}^{\mathbf{P}}_{\mathbf{F2,i}}}{\bar{W}_i}$	-0.00043* (0.00022)	0.00006 (0.00075)	-0.00057 (0.00080)
Fixed Hiring Cost Temp:	$\frac{\boldsymbol{\theta}^{\mathbf{T}}_{\mathbf{H0,i}}}{\bar{W}_i}$	**1.417**** **(0.060)**	0.097** (0.039)	0.049 (0.046)
Linear Hiring Cost Temp:	$\frac{\boldsymbol{\theta}^{\mathbf{T}}_{\mathbf{H1,i}}}{\bar{W}_i}$	**0.181**** **(0.049)**	**0.107**** **(0.041)**	**0.089**** **(0.045)**
Quad Hiring Cost Temp:	$\frac{\boldsymbol{\theta}^{\mathbf{T}}_{\mathbf{H2,i}}}{\bar{W}_i}$	0.00008 (0.00041)	-0.00007 (0.00059)	-0.00081 (0.00086)
Fixed Firing Cost Temp:	$\frac{\boldsymbol{\theta}^{\mathbf{T}}_{\mathbf{F0,i}}}{\bar{W}_i}$	0.067 (0.094)	0.061 (0.084)	0.058 (0.113)
Linear Firing Cost Temp:	$\frac{\boldsymbol{\theta}^{\mathbf{T}}_{\mathbf{F1,i}}}{\bar{W}_i}$	**0.098**** **(0.045)**	**0.060 (0.037)**	**0.051 (0.048)**
Quad Firing Cost Temp:	$\frac{\boldsymbol{\theta}^{\mathbf{T}}_{\mathbf{F2,i}}}{\bar{W}_i}$	-0.00006 (0.00106)	0.00037 (0.00085)	0.00055 (0.00092)
	$\frac{\boldsymbol{\sigma}_{P,i}}{\bar{W}_i}$	0.611 (0.080)	0.515 (0.058)	0.540 (0.079)
	$\frac{\boldsymbol{\sigma}_{T,i}}{\bar{W}_i}$	0.170 (0.025)	0.172 (0.019)	0.149 (0.026)
	$\frac{\boldsymbol{\sigma}_{0,i}}{\bar{W}_i}$	0.871 (0.069)	0.710 (0.054)	0.761 (0.064)
# Observations		2,274	7,219	6,257
Likelihood Ratio Index[2]		0.232	0.220	0.267

Note 1: All the parameters are unit-free because all firing and hiring costs are proportional to the firm-specific mean wage $\bar{W}_i$.
Note 2: The Likelihood Ratio Index is a measure of goodness of fit that is defined as $1 - (\log \hat{L} / \log L_0)$, where $\log \hat{L}$ is the log-likelihood of the estimated model, and $\log L_0$ is the log-likelihoo under the the hypothesis that all parameters except $\boldsymbol{\sigma}_{0,i}$ are equal to zero.

Table 6
Goodness of Fit Measures of the Estimated Model
Unbalanced panel 1982-1993 (2356 firms)

Statistics	**Period 1983-1984 Model (Empirical)**	**Period 1985-1988 Model (Empirical)**	**Period 1989-1992 Model (Empirical)**
Permanent Employment per Firm (Median)	98.0 (95.0)	64.0 (66.0)	59.0 (56.0)
Proportion of Temporary Workers (Mean)	4.4% (4.3%)	6.6% (6.9%)	11.8% (11.3)
Percentage of Zeroes in ΔL^P	14.8% (15.1%)	18.1% (18.8%)	19.7% (19.6%)
Percentage of Zeroes in ΔL^T	46.9% (52.8%)	39.8% (43.9%)	28.1% (32.5%)
Median Value of d^P Conditional of $d^P > 0$	4.0% (3.9%)	4.8% (5.2%)	5.6% (5.7%)
Median Value of d^P Conditional of $d^P < 0$	-3.9% (-3.8%)	-4.0% (-4.3%)	-5.1% (-5.2%)
Median Value of d^T Conditional of $d^T > 0$	1.7% (1.7%)	2.9% (2.7%)	4.0% (4.2%)
Median Value of d^T Conditional of $d^T < 0$	-1.3% (-1.4%)	-2.3% (-2.0%)	-3.7% (-3.8%)
Cross-sectional Variance log Perm. Employment	1.64 (1.66)	1.66 (1.72)	1.56 (1.59)
Cross-sectional Variance log Total Employment	1.54 (1.59)	1.60 (1.64)	1.49 (1.51)

Table 7
Evaluation of the Labor Market Reform

Statistics	Pre-Reform Economy$^{(1)}$	Post-Reform Economy$^{(1)}$	Counterfactual Reform$^{(1)}$
Permanent Employment per Firm (Median)	99.0	89.1 (-10.0%)	105.5 (+6.6%)
Total Employment per Firm (Median)	102.7	106.3 (+3.5%)	106.9 (+4.1)
Proportion of Temporary Workers (Mean)	3.8%	16.2%	1.3%
Median Absolute Value of d^P	2.9%	2.5%	4.2%
Median Absolute Value of d^T	0.0%	3.1%	0.9%
Output per Firm (Median)$^{(2)}$	100	100.7	101.9
Value of a Firm (Median)$^{(2)}$	100	101.2	104.8

Note 1: The values of the structural parameters are: for the pre-reform model, the ones estimated for the period 1983-1984; for the post-reform model, the ones estimated for the period 1989-1992; for the counterfactual model, we consider the 1983-1984 parameters except for the linear firing costs θ^P_{F1} and θ^T_{F1} which reduced by half.
Note 2: We normalize at 100 output-per-firm and value of a firm in the pre-reform model.

Figure 1: Unemployment rate and GDP growth in Spain.

Source: Spain's Labor Force Survey and National Accounts.

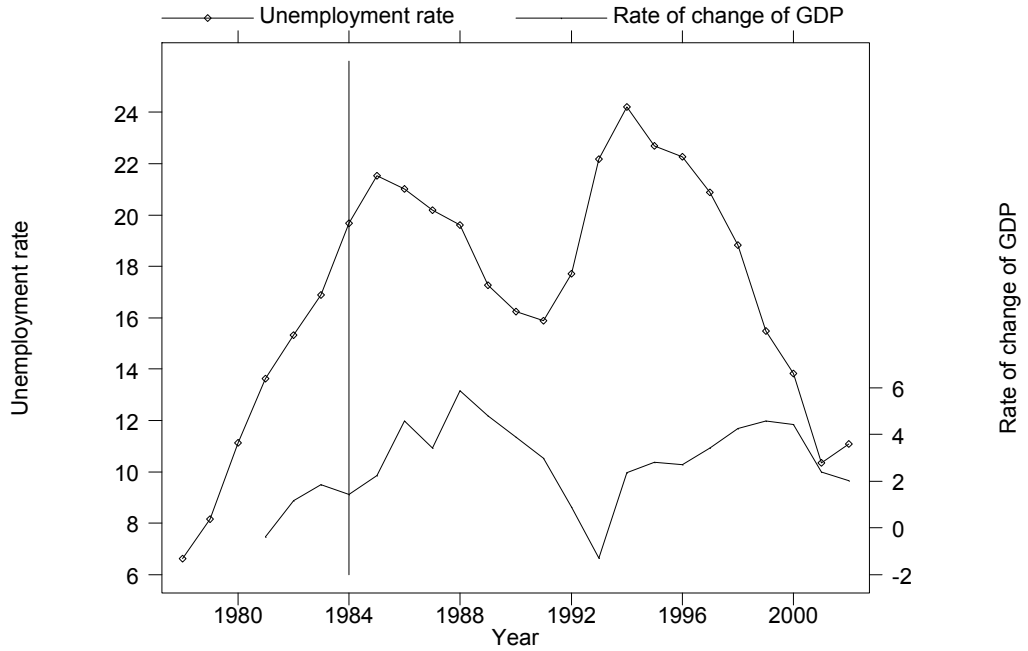


Figure 2: Share of temporary employment in total employment.

Source: Spanish Labor Force Survey and CBBE.

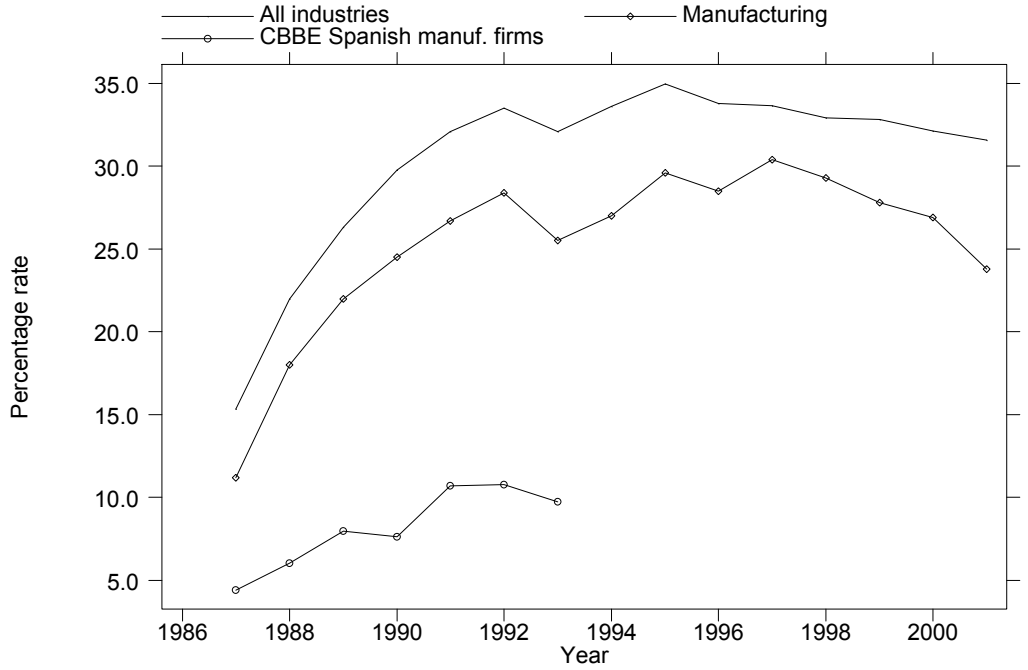


Figure 3: Rates of growth of output and employment.

Source: CBBE sample of Spanish manufacturing firms.

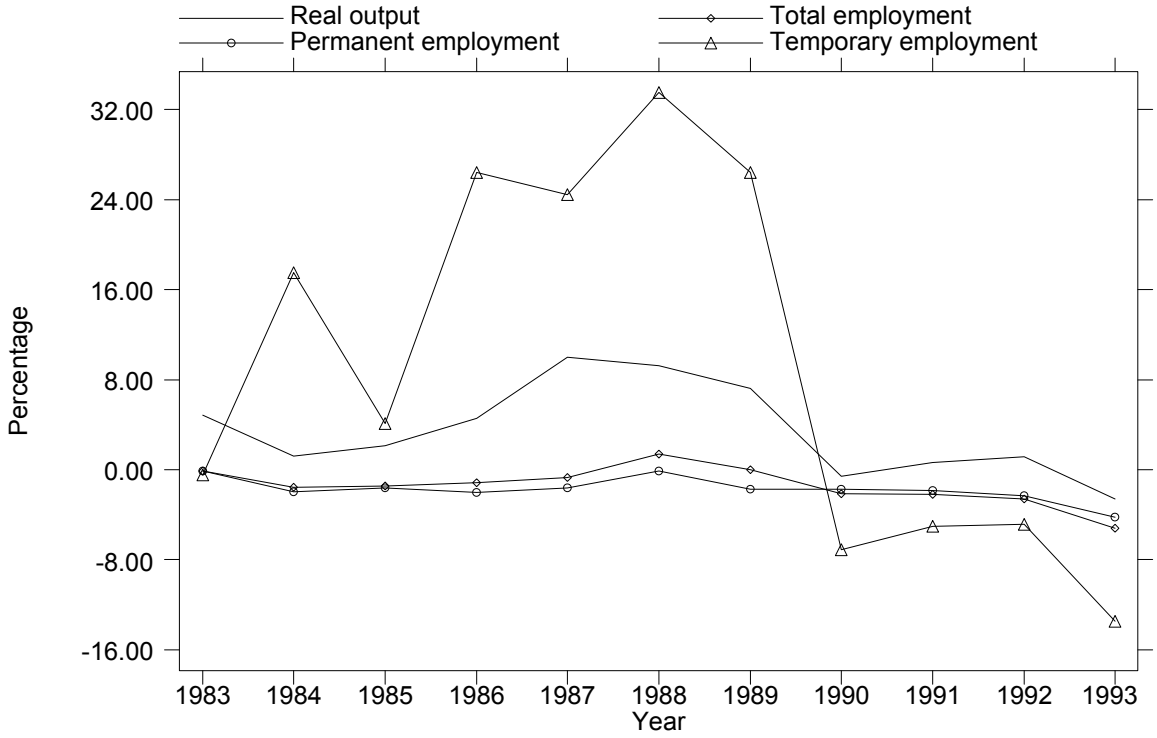


Figure 4: Share of temporary employment in total employment, by firm average size.

Source: CBBE sample of Spanish manufacturing firms.

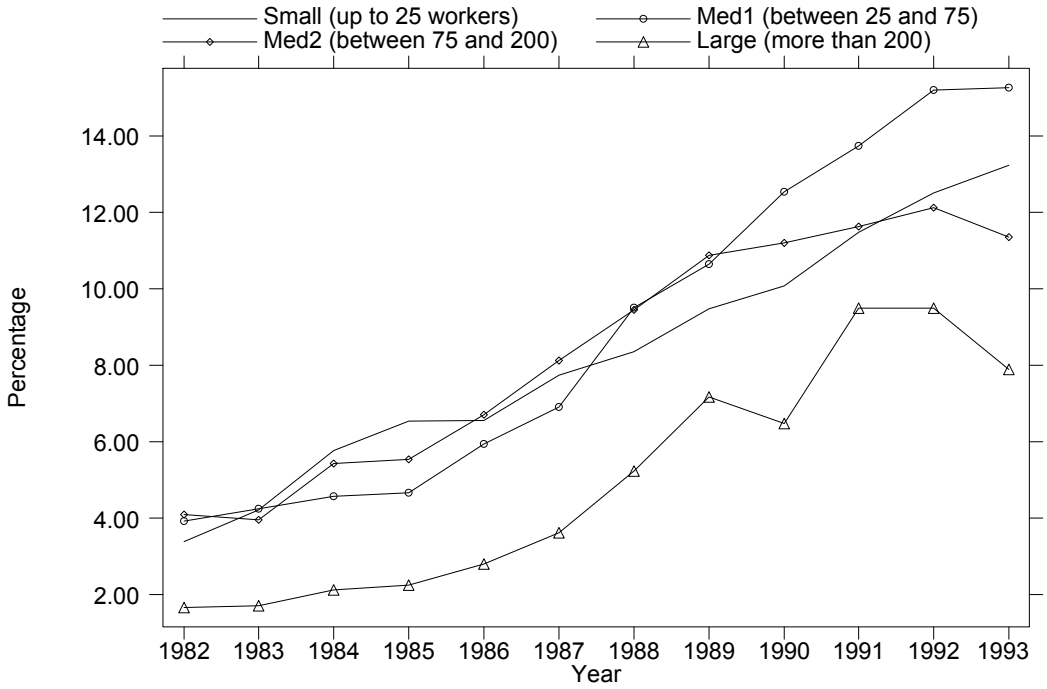


Figure 5: Rates of job creation and job destruction by type of contract (weighted averages).

Source: CBBE sample of Spanish manufacturing firms.

(The rates for job destruction appear with negative sign)

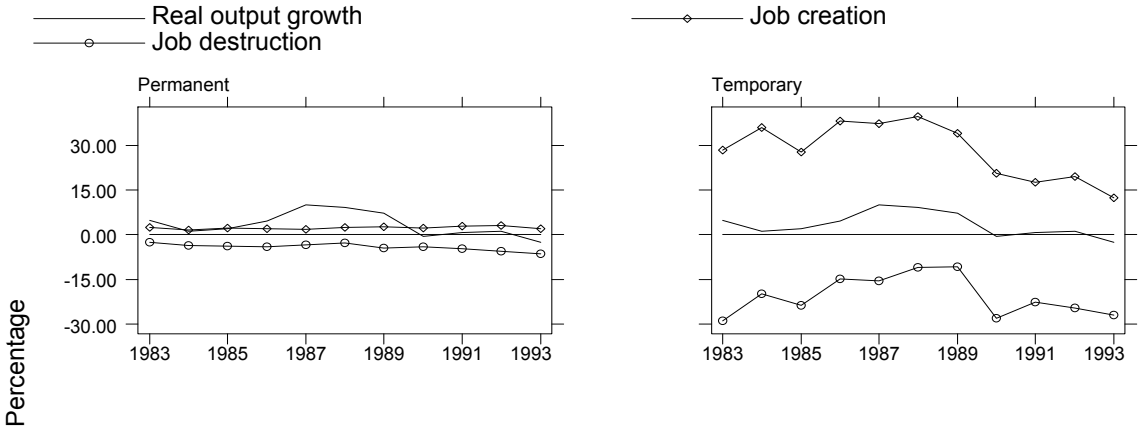


Figure 6: Net changes in permanent employment.

Source: CBBE sample of Spanish manufacturing firms

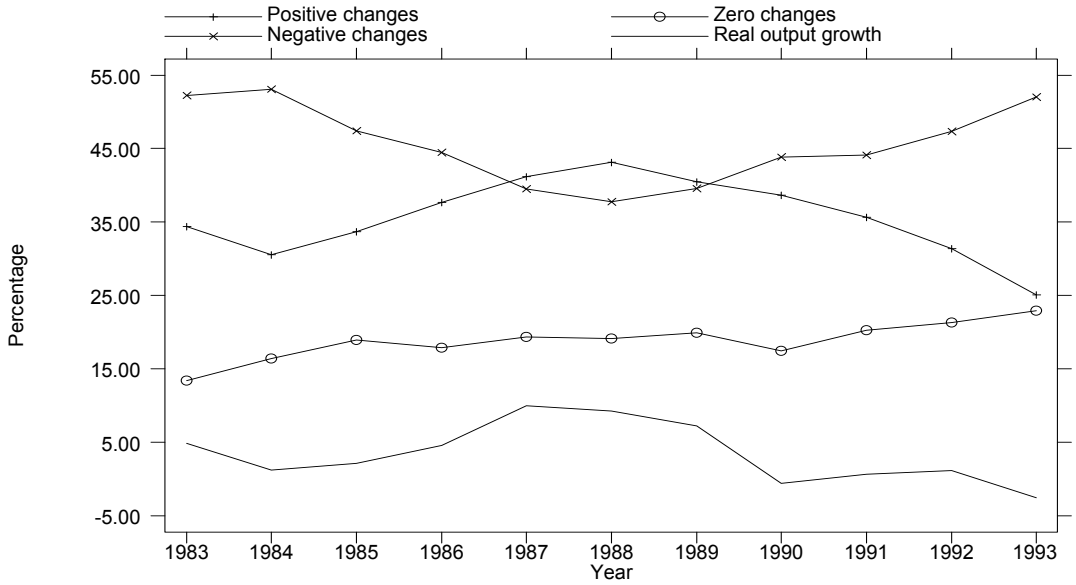


Figure 7: Time Series of the Estimated Average Wages

of Permanent and Temporary Workers

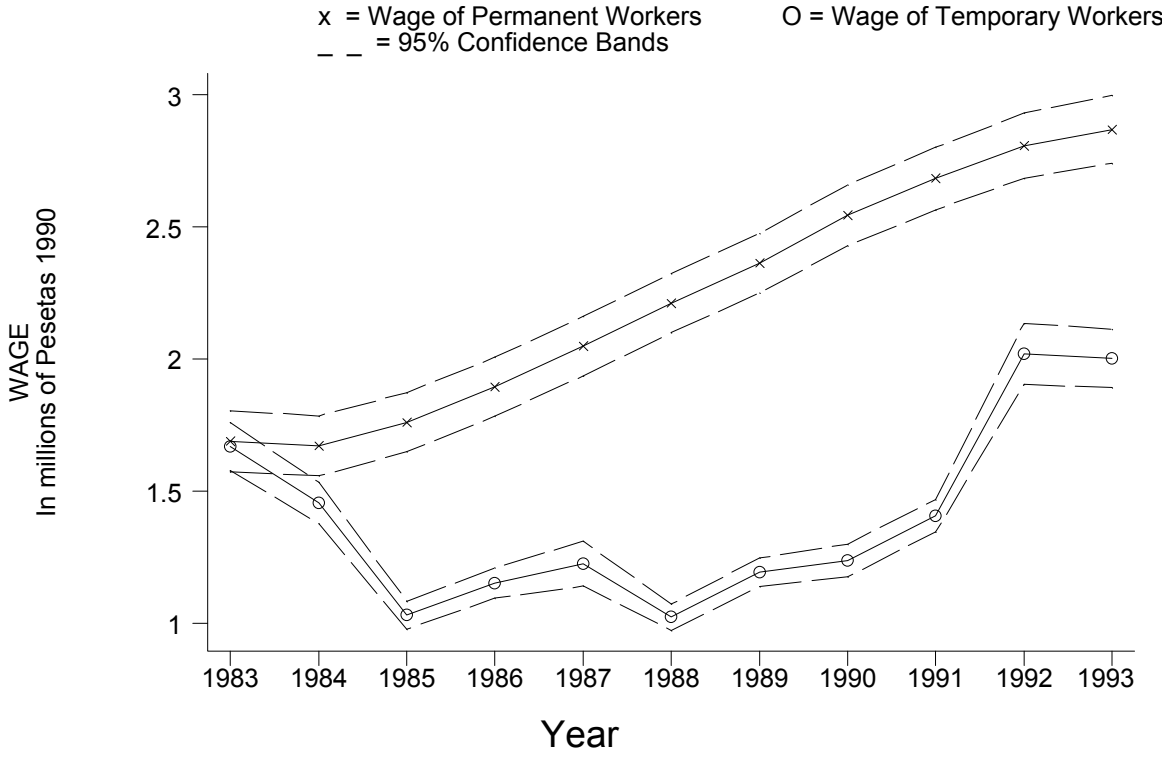
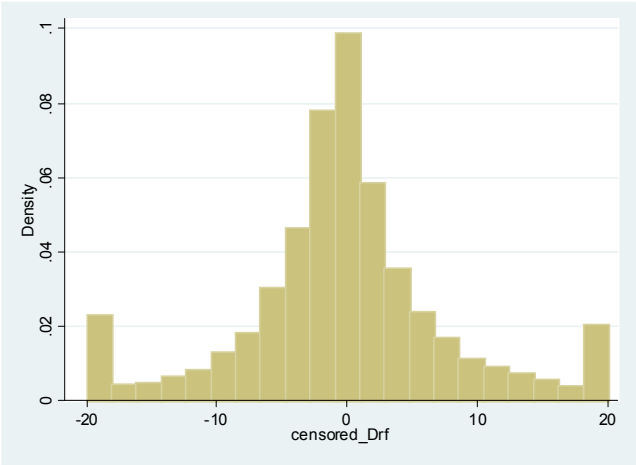
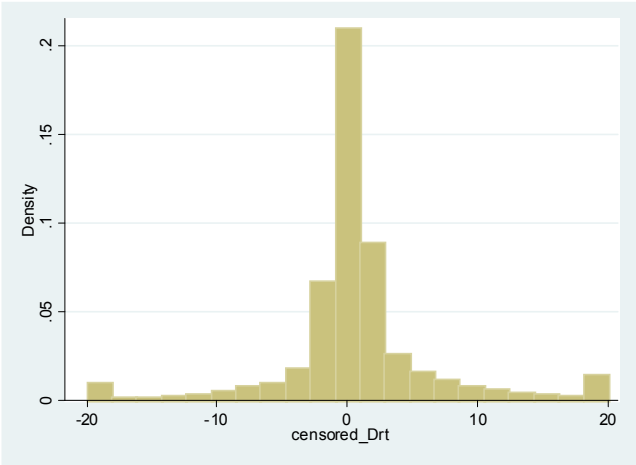


Figure 8: Histograms of Discretized Decision and State Variables

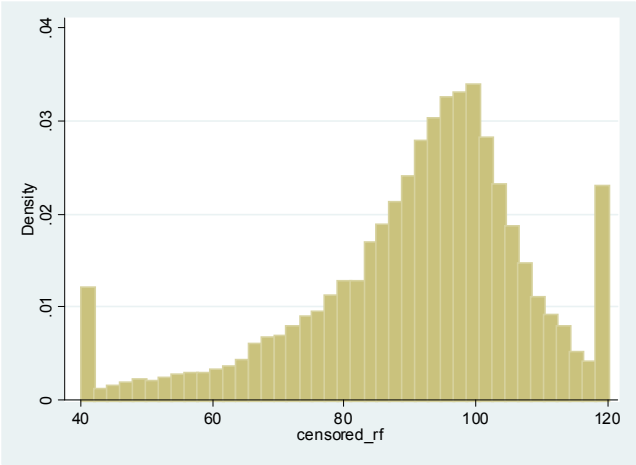
Change in Permanent Employment (d^P)



Change in Temporary Employment (d^T)



Level of Permanent Employment (ℓ^P)



Level of Temporary Employment (ℓ^T)