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LABOR CONTRACTS AND FLEXIBILITY: EVIDENCE FROM A LABOR MARKET REFORM IN SPAIN

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This paper evaluates the effects of a labor market reform in Spain that removed restrictions on fixed-term or temporary contracts. Our empirical results are based on longitudinal firm-level data that cover observations before and after the reform. We posit 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 firms are very small. In contrast, a counterfactual reform that halved all firing costs would produce the same employment increase as the actual reform, but much larger improvements in productivity and in the value of firms. (*JEL* J23, J32, J41)

I. 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. The consequences of job security provisions on labor market performance have been broadly analyzed both at the theoretical and at the empirical level. From a theoretical point of view, the effect of firing costs on employment is ambiguous. Firing costs reduce both hiring during expansions, and dismissals during downturns. The net effect depends on different factors, including the size of hiring and firing costs and the persistence of demand and supply shocks. The empirical 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, particularly regarding the effects on the level of employment. Therefore, the employment effect of firing costs is an empirical question that should be evaluated case by case. The labor market reforms implemented by several European countries since the 1980s provide unique information to identify the effect of firing costs on labor market outcomes.

In this article, we study the effects on employment, job turnover, and firms' productivity of a Spanish labor market reform that took place in 1984, which liberalized the use of temporary contracts and reduced the redundancy payments at termination of these contracts. After the reform, temporary contracts could be applied

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ABBREVIATIONS

CBBE: Balance Sheets of the Bank of Spain
ET: Workers' Statute
GDP: Gross Domestic Product
MLE: Maximum Likelihood Estimator
NPL: Nested Pseudo Likelihood
OECD: Organization for Economic Co-operation and Development

by any firm, irrespective of their size, industry, or performance, to any type of worker, irrespective of their occupation, age, or gender. Nevertheless, the stringent dismissal regulations for indefinite-duration or permanent contracts remained unchanged after the reform.

Our approach is based on the estimation of a micro-econometric dynamic structural model of labor demand, with permanent and temporary employment, and on the comparison between the pre-reform and post-reform periods in the estimates of the structural parameters and in the steady-state distributions of employment. Our primary source of data consists of a longitudinal sample of 2,356 Spanish manufacturing firms, during the period 1982–1993, with information on employment by type of contract, capital, output, and wages. In our model, we consider that the reform may introduce changes in firing and hiring costs of temporary workers, and in their productivity relative to permanent workers.

The estimated hiring and firing costs for temporary workers are significantly lower after the reform. Experiments using the estimated model show important positive effects of the reform on total employment (i.e., a 3.5% increase) and job turnover in steady state. Further, the reform led to a strong substitution of permanent by temporary workers, with a 10% decline in permanent employment. The effect on labor productivity is negative, and the effect on the value of firms is negligible.

We also compare the actual reform with a counterfactual reform consisting of halving firing costs for both types of contracts. Although the employment increases under both reforms are alike, the employment composition by contract is very different. As a consequence, such counterfactual reform had led to much larger improvements in the productivity and the value of firms. Compared with this counterfactual reform, the factual introduction of temporary contracts leads to excess turnover and a large proportion of employment of workers with low firm-specific experience.

The structural approach is justified on several grounds. 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 differently 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 papers on dynamic structural models of labor demand with nonconvex adjustment costs, such as Rota (2004), and Cooper and Willis (2004). The most relevant extensions are the following: (a) we consider two types of labor contracts, temporary and permanent; (b) our specification of labor adjustment costs is very general and allows for fixed, linear, and quadratic, asymmetric adjustment costs which can be different for the two types of contracts; and (c) 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.

As our dataset only provides information on employment stocks, but not on employment flows, our specification for labor adjustment costs is defined in terms of net employment changes. The lack of information about gross employment changes prevents to disentangle among quantitatively similar net employment changes which result from simultaneous hires, layoffs, and voluntary quits. Several papers using datasets that directly measure gross employment flows, like Abowd, Corbel, and Kramarz (1999) and Goux, Maurin, and Pauchet (2001), have shown how net employment changes can result from large simultaneous flows of hiring, firing, and voluntary quits. In particular, distinguishing between layoffs and voluntary quits can make a substantial difference in the case of permanent workers, with high severance payments. In an earlier version of this article, Aguirregabiria and Alonso-Borrego (1999) exploited complementary information on severance payments to distinguish between costly and noncostly reductions in net employment, obtaining firing cost estimates that are larger than the ones that ignore

1. See Pfann and Palm (1993), Hamermesh (1993), and the survey paper by Bond and Van Reenen (2007) for references.

this distinction.[2] The employment effects of the labor market reform using those estimates of firing costs are stronger than the ones that we report in this paper.

The database that we exploit in this paper over-samples large firms. However, we believe that our main results on the effects of the labor market reform are even stronger for the whole population of Spanish firms. In particular, the estimation of the model shows that the effects of the reform on the level of employment, proportion of temporary labor, and job turnover is significantly larger for smaller firms.

The rest of the paper is organized as follows. In Section II, 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 III describes our dataset. In Section IV, 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 V. We present experiments that evaluate the effects of the reform in Section VI. Section VII summarizes our main findings and concludes.

II. THE ROLE OF JOB SECURITY PROVISIONS

A. *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 Organization for Economic Co-operation and Development (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 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, 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 (1992) finds that job security provisions

2. When exploiting information on severance payments, we defined negative employment changes as layoffs if severance payments were strictly positive; otherwise, we imputed them as separations due to voluntary quits. Under such definitions, a large fraction of negative employment changes were imputed as voluntary quits, and therefore we measured a frequency and an amount of layoffs which were substantially smaller. Still, we were not able to identify simultaneous hires, layoffs, and voluntary quits. However, we can reasonably argue that the frequency and amount of simultaneous hires and layoffs is lower than the frequency of simultaneous layoffs and voluntary quits. Therefore, we believe that firing cost estimates based on net employment changes that ignore voluntary quits, as well as simultaneous hires and layoffs, provide a lower bound to the firing cost parameter that we would have obtained if we observed the different gross employment flows.

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 U.S. 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 in Spain 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. Blanchard and Landier (2002) and Cahuc and Postel-Vinay (2002) consider matching models to assess the effect of partial labor market reforms like the one undertaken in Spain in 1984, finding that although fixed-term contracts tend to foster job creation, the impact on employment might be negative. Their results also suggest that narrowing firing costs differences between permanent and fixed term would increase employment. In a similar strand, Costain, Jimeno, and Thomas (2010) study the cyclical properties of the Spanish dual labor market, looking both at unemployment level and unemployment volatility. Interestingly, they find that to eliminate duality without raising the level of unemployment, firing costs should be substantially cut. Bentolila et al. (2012) consider France and Spain, which share similar labor market institutions except for a larger gap in firing costs between permanent and temporary workers, and laxer rules on the use of temporary contracts in Spain than in France. Their calibration results show that these differences account for the strikingly larger increase in unemployment during the current Great Recession in Spain with respect to France. Güell and Rodríguez-Mora (2010) analyze the effects of introducing temporary contracts in the context of an efficiency wage model, concluding that temporary contracts may increase unemployment if rigidities on wage setting, such as minimum wages, exist. Álvarez and Veracierto (2012) extend an Islands model with undirected search and complete markets to deal with severance taxes conditional on tenure. They interpret this dependence as a form of temporary contracts. In a similar framework, Veracierto (2001) assesses the short-run consequences of introducing labor market flexibility. Both papers find that fixed-term contracts may increase unemployment.

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 have affected labor market outcomes using a differences-in-differences approach. Kugler (2004) exploits the variation in employment protection between workers covered and not covered by the legislation, to study the effect of firing costs on unemployment hazard, using Colombian individual data before and after a reform which reduced firing costs. In a similar fashion, Kugler and Pica (2008) exploit the differences in employment protection among small and larger firms to study the effect on worker and job flows of firing costs. The results provide evidence about the negative effect of severance payments on both employment level and employment changes. Using establishment-level data for the United States and the variation across states in dismissal costs, Autor, Kerr, and Kugler (2007) find that dismissal costs reduce employment flows and total factor productivity. 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 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, which dates back to the seminal paper by Sargent (1978). As mentioned in Section I, our model builds on and extends the recent literature on dynamic structural models of labor demand with nonconvex adjustment costs (Rota 2004; 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.

B. 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.[3] 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. 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.

Until the late 1970s, 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 duration. This institutional framework 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's wages) if the dismissal is considered "fair," and 45 days (up to a maximum of 42 months' wages) if it is considered "unfair." There are two reasons for fair dismissals: those attributable to the worker (incompetence or negligence in his job), and other objective reasons that cannot be attributed to the worker (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 acknowledge the dismissal—as it is usually the case—he may appeal to the labor court. This obliges the firm to bear a legal process, during which it should pay interim wages while labor courts reach their decisions (up to 2 months). Given that labor courts are in many cases favorable to the workers, the agreed severance payments can even exceed the statutory amounts for unfair dismissals. Besides, any dismissal (except for disciplinary reasons) requires written advance notice of 30 days. These job security rules for *permanent* workers remained unchanged until 1997.[4]

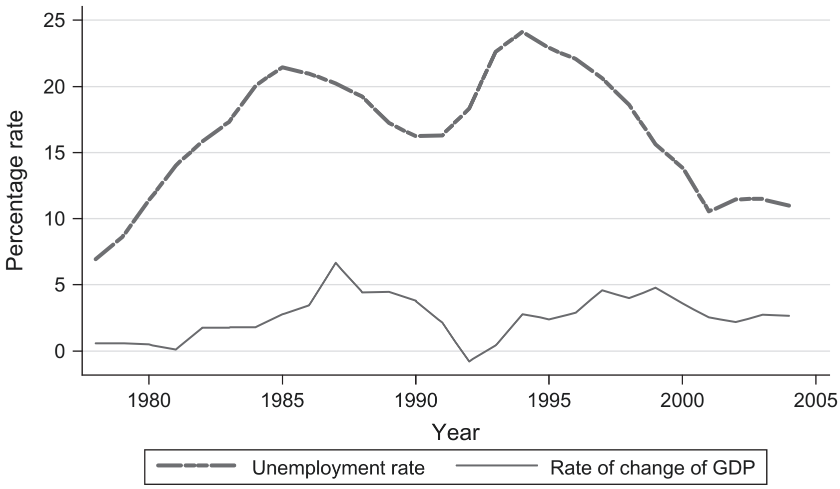
The ET allowed *temporary* or fixed-term contracts, which could be cancelled at termination with a much smaller severance payment and without court or regulatory intervention. However, their use was mainly limited to jobs that were temporary in nature because of the seasonal nature of the production activity, the need to cover absent posts, or the start-up of a new firm. These contracts had been used earlier to a certain extent in agriculture, construction, and services industries; however, their use in manufacturing had been very scarce.

Figure 1 shows the evolution of gross domestic product (GDP) growth and unemployment in Spain. Despite the fact that GDP was growing at the beginning of the 1980s, the unemployment rate kept rising, and by the end of 1984, it 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 noncausal fixed-term contracts. The main feature of the reform was that the use

3. 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.

4. 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 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 its scope was limited to certain types of workers.

FIGURE 1
Unemployment Rate and GDP Growth in Spain



Source: Spanish Labor Force Survey and National Accounts.

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 (3, 6, or 12 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, the reform established that temporary contracts could only be renewed up to 3 years, when the firm should decide whether to offer the worker a permanent contract or to dismiss him.[5] 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 soared. Spain became by far the European country with the highest percentage of temporary employment, which amounted to 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.

III. DATA AND PRELIMINARY EVIDENCE

The main dataset consists of an unbalanced panel of 2, 356 nonenergy manufacturing firms collected from the database of *Central de Balances del Banco de España* (Balance Sheets of the Bank of Spain, CBBE hereinafter) between 1982 and 1993. The firms included in the raw database represent almost 40% of the total value added in Spanish manufacturing. This dataset provides 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. Due to the fact that response is completely voluntary, the largest firms are overrepresented in the sample. Therefore, care must be taken in extrapolating to the whole population of Spanish firms the results based on our sample. The criteria for selection of the sample and construction of the variables used in the empirical analysis (market value of the capital stocks, wages, etc.) are described in Appendix A.

Wage information by type of contract is available from the Spanish wage distribution survey (*Distribución Salarial*). However, that survey is only available for the years 1988 and 1992,

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

and it provides only aggregate information. We describe in Section IV.A our approach to obtain estimates of wages of permanent and temporary workers for the whole period 1982–1993.

Figure 2 presents the evolution of the proportion of temporary workers in the Spanish economy and in our sample.[6] The share of temporary contracts, which was estimated to be about 10% of total employment and 3% of manufacturing employment in 1984, rose to 35% and 31%, respectively, in 1995, remaining at high levels since then. It is not surprising that the share of temporary employment increased gradually since the reform. Due to the substantial firing costs for permanent workers, firms could wait until some permanent workers retire or quit voluntarily to circumvent severance payments. Figure 2 shows also a large disparity between the proportion of temporary workers from the Labor Force Survey and the proportion from the CBBE. The main factor to explain this discrepancy is that CBBE overrepresents large manufacturing firms, and this type of firm tends to have a smaller proportion of temporary workers (see Figure 4 below).

In Figure 3, we report the evolution of the growth rates of real output and employment for our sample. The sample period covers an expansion, 1986–1989, and a recession, 1990–1993. However, both the number and the proportion of permanent employees exhibit a monotonous decrease. After the introduction of the reform, in November 1984, temporary employment rose significantly from 1986 to 1990 and decreased from 1990 to 1993, and its share in total employment rose from 2.89% in 1985 to 9.72% in 1993. 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 manufacturing (in fact, the correlation coefficient between both series is above 90%). However, the figure for our sample is substantially below the aggregate figures, which were well above 20% at the beginning of the 1990s. This discrepancy is due to the fact that larger companies, which are overrepresented in our sample, are more prompted to use permanent employment than small or medium ones. In Figure 4, we can see that in our sample the proportion of temporary employment decreases with size.

6. Unfortunately, the Spanish Labor Force Survey did not report any information about the type of contract before 1987.

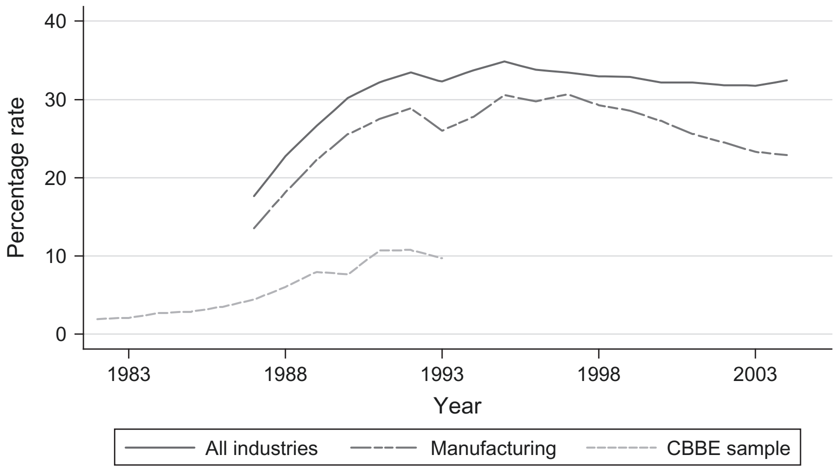
Figure 5 presents the job creation and job destruction rates for permanent and temporary employment using the statistics proposed by Davis and Haltiwanger (1992).[7] The small job turnover rates for permanent employment contrasts with the very high rates for temporary employment. Furthermore, the creation 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 evolution in the proportion of our 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 1 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 389 firms in the two periods. The post-reform period has been also selected for comparability reasons, such that the pattern of output growth is very similar to the pre-reform period. From this comparison, we can establish some interesting facts. The median employment growth is similar in the two periods, but it is significantly more dispersed after the reform. The shrink in permanent employment is compensated by temporary employment, so that total employment exhibits an increase of 5.4% for the median firm in terms of number of workers, 8.3% for smaller firms, and 0.3% for larger firms. We can see that firm productivity, as measured by the sales to wage bill ratio, goes down after the reform. Interestingly, in Section VI, our counterfactual experiments based on the estimated model show similar magnitudes for the steady-state effects of the labor market reform on total employment,

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

FIGURE 2
Share of Temporary Employment in Total Employment



Note: Spanish Labor Force Survey did not report information about the type of contract before 1987.
Source: Spanish Labor Force Survey and CBBE sample of Spanish manufacturing firms.

proportion of temporary workers, and productivity.

IV. MODEL

A. 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. 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.

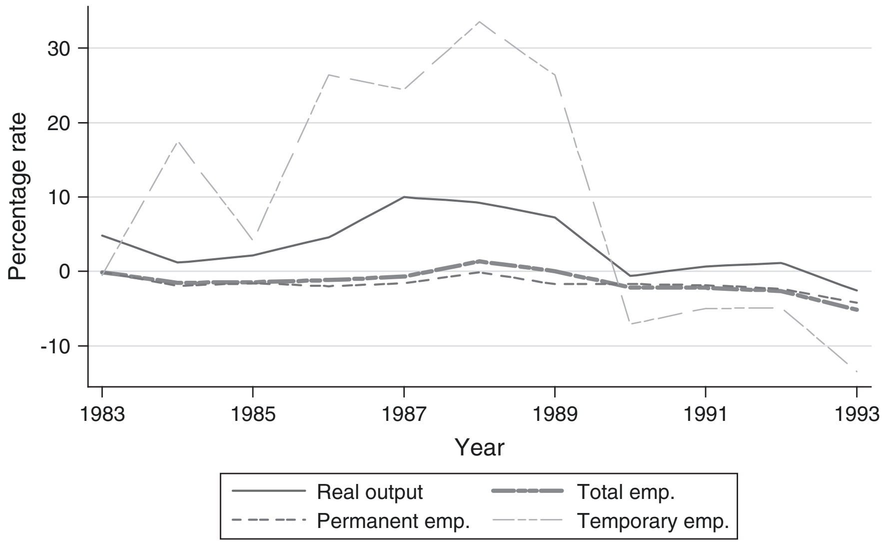
Firms produce a homogeneous good using labor as the only variable input, and sell their output in a competitive market.[8] Every period t, the firm chooses the amounts of permanent and temporary labor that maximize its expected intertemporal profit, $\mathbb{E}_t\left(\sum_{j=0}^{\infty}\beta^j\Pi_{t+j}\right)$, where $\mathbb{E}_t$ is the conditional expectation 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:

$$\Pi_t = Y_t - \mathrm{WB}_t - \mathrm{AC}_t + \xi_t, \tag{1}$$

where Y_t is real output, $\mathrm{W}B_t$ is the wage bill, $\mathrm{A}C_t$ 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.

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

FIGURE 3
Rates of Growth of Output and Employment



Source: CBBE sample of Spanish manufacturing firms.

The production technology is described by the production function

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

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

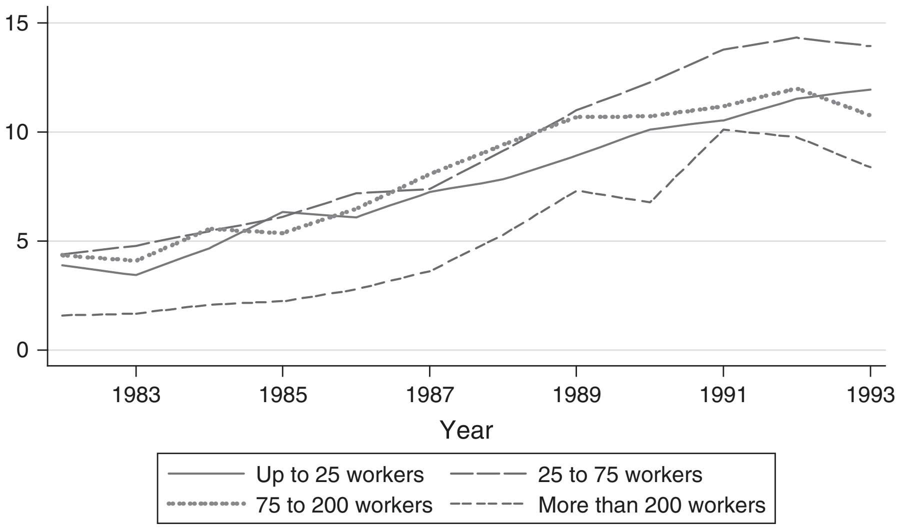
As mentioned above, dismissal cost is the only exogenous difference between temporary and permanent workers. Therefore, if we had the same type of workers under each contract, they would be perfect substitutes with one-to-one rate of substitution. However, as we have mentioned earlier, differences in dismissal costs can endogenously generate differences in the relative productivity of permanent and temporary workers. We account for such differences through the parameter λ, which measures the productivity of temporary workers with respect to permanent workers. We could have otherwise considered a constant elasticity of substitution technology of production in permanent and temporary workers to allow certain complementarity between worker types. However, in this article we do not see permanent and temporary workers as different labor inputs. Besides, the existence of labor adjustment costs and, importantly, uncertainty about future productivity and wages, implies that, despite the perfect substitution between worker types, the optimal proportion of temporary employment is not either zero or one in most of the cases.

The wage bill is $WB_t = W_t^{\mathrm{T}} L_t^{\mathrm{T}} + W_t^{\mathrm{P}} L_t^{\mathrm{P}}$, 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 wages $W_t = (W_t^{\mathrm{T}}, W_t^{\mathrm{P}})$ follows a first-order Markov process with transition probability function $f_W(W_{t+1}|W_t)$.

Our sample information on the firm's total wage bill is not broken down by type of contract. The assumption that the wage of temporary workers is the same across firms is a restriction that we use to identify the evolution over time in the average wage ratio between temporary and permanent workers. We describe this identification approach below. Temporary workers do not have bargaining power or firm-specific experience. Empirical evidence from other data sources with information on wages at the worker level (Wage Structure Survey 1995) shows that, though the wages of temporary

FIGURE 4

Share of Temporary Employment in Total Employment by Firm Average Size



Source: CBBE sample of Spanish manufacturing firms.

workers varies across firms, they have a much weaker correlation with measures of firm size than the wages of permanent workers. An implication of our assumption is that we interpret firms with high wage-bill-to-permanent-workers ratio as firms with costly permanent workers, such that we expect that, *ceteris paribus*, these firms are more willing to substitute permanent workers with temporary workers.

The wage bill of firm i at year t is $\mathrm{W}B_{it} = W_{it}^{\mathrm{T}} L_{it}^{\mathrm{T}} + W_{it}^{\mathrm{P}} L_{it}^{\mathrm{P}}$. Given the assumption that the wage of temporary workers is the same for every firm, we have that:

$$(3) \qquad WB_{it}/L_{it}^{\mathrm{P}} = W_t^{\mathrm{T}} \left(L_{it}^{\mathrm{T}}/L_{it}^{\mathrm{P}}\right) + W_{it}^{\mathrm{P}}.$$

We observe $\mathrm{W}B_{it}/L_{it}^{\mathrm{P}}$ and $L_{it}^{\mathrm{T}}/L_{it}^{\mathrm{P}}$ but we do not have data on W_t^{T} and W_{it}^{P}. If the wage of permanent workers were mean independent of the temporary-to-permanent ratio, $L_{it}^{\mathrm{T}}/L_{it}^{\mathrm{P}}$, we could estimate the value W_t^{T} by running an OLS regression for $\mathrm{W}B_{it}/L_{it}^{\mathrm{P}}$ on $L_{it}^{\mathrm{T}}/L_{it}^{\mathrm{P}}$ at year t. 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:

$$(4) \qquad W_{it}^{\mathrm{P}} = \mu_i + \gamma_t + u_{it}$$

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}.

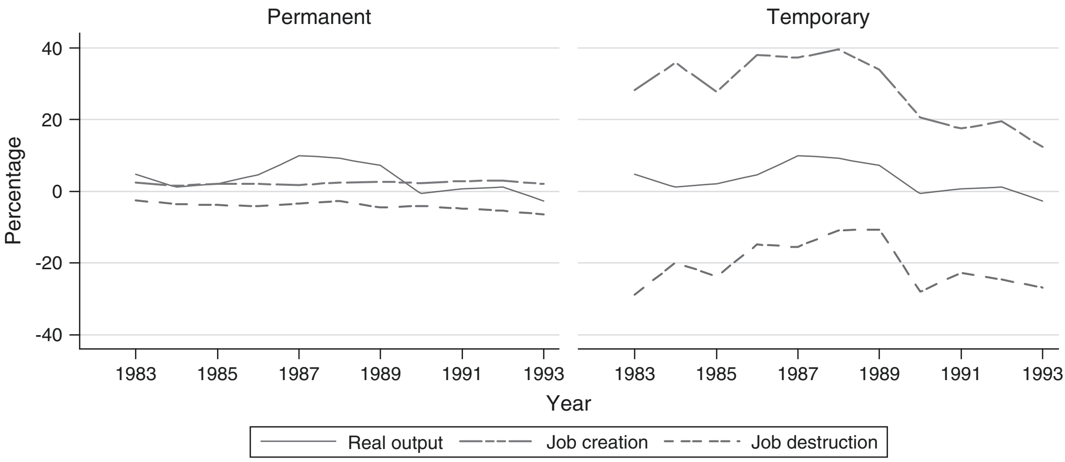
The specification of labor adjustment costs, defined in terms of net employment changes, for each type of employment, $\Delta L_t^j \equiv L_t^j - L_{t-1}^j$ ($j = \mathrm{T,P}$), includes both lump-sum and linear components:

(5)

$$\begin{aligned} AC_t = {} & 1(\Delta L_t^{\mathrm{T}} > 0) \\ & \times \left[\theta_{H0}^{\mathrm{T}} + \theta_{H1}^{\mathrm{T}} \Delta L_t^{\mathrm{T}} + \theta_{H2}^{\mathrm{T}} \left(\Delta L_t^{\mathrm{T}}\right)^2\right] \\ & + 1(\Delta L_t^{\mathrm{T}} < 0) \left[\theta_{F0}^{\mathrm{T}} - \theta_{F1}^{\mathrm{T}} \Delta L_t^{\mathrm{T}} + \theta_{F2}^{\mathrm{T}} \left(\Delta L_t^{\mathrm{T}}\right)^2\right] \\ & + 1(\Delta L_t^{\mathrm{P}} > 0) \left[\theta_{H0}^{\mathrm{P}} + \theta_{H1}^{\mathrm{P}} \Delta L_t^{\mathrm{P}} + \theta_{H2}^{\mathrm{P}} \left(\Delta L_t^{\mathrm{P}}\right)^2\right] \\ & + 1(\Delta L_t^{\mathrm{P}} < 0) \left[\theta_{F0}^{\mathrm{P}} - \theta_{F1}^{\mathrm{P}} \Delta L_t^{\mathrm{P}} + \theta_{F2}^{\mathrm{P}} \left(\Delta L_t^{\mathrm{P}}\right)^2\right] \end{aligned}$$

where $1(.)$ is the binary indicator function; $\{\theta_{H0}^j, \theta_{H1}^j, \theta_{H2}^j, \theta_{F0}^j, \theta_{F1}^j, \theta_{F2}^j : j = \mathrm{T,P}\}$ are (nonnegative) parameters. The first two summands refer

FIGURE 5
Rates of Job Creation and Job Destruction by Contract Type



Note: The rates for job destruction appear with negative sign.
Source: CBBE sample of Spanish manufacturing firms.

to hiring and firing costs of temporary workers. The third and fourth terms are the corresponding adjustment costs for hiring and firing permanent workers.

There are different sources associated with adjustment of labor inputs, the most obvious being those associated to technological disruption and, very especially, the legal costs imposed by the regulations on hiring and firing. But there are further sources that can magnify or mitigate the total costs of adjustment. In particular, in the case of firing, the firm can bear additional costs if the workers go to the labor court adducing unfair dismissal: litigation costs, and the costs associated with the firms obligation to pay the wage of the fired worker until the case is settled by the court (see Galdón-Sánchez and Güell 2000). On the contrary, firms and dismissed workers may eventually settle an agreement to either avoid litigation by which dismissal compensation may be below the legal redundancy payment for unfair dismissal, or pay the worker outplacement services, or negotiate payment for early retirement. In fact, we are not imposing that adjustment costs are necessarily above or below the administrative ones, using a revealed preference approach.

For the sake of simplicity, and based on the empirical evidence, the model does not account explicitly for the firm's decision on conversion of temporary contracts into permanent contracts. Dolado, García-Serrano, and Jimeno (2002) and Güell and Petrongolo (2007) indicate that the conversion rate of temporary into permanent contracts in Spain is low, reflecting the fact that "employers use those contracts more as a flexible device to adjust employment in the face of adverse shocks than as a screening device."

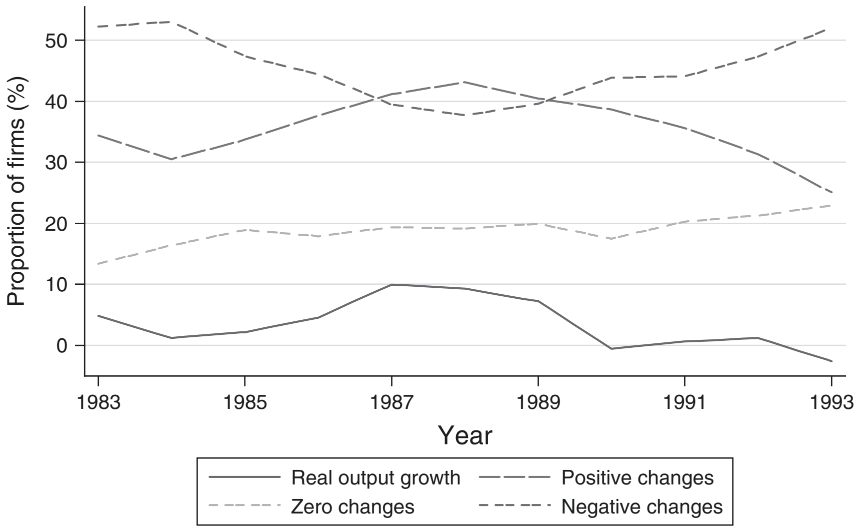
The firm chooses employment changes so as to maximize its expected intertemporal profit. We consider a discrete choice model such that the set of possible values of $(\Delta L_t^{\mathrm{P}}, \Delta L_t^{\mathrm{T}})$ is discrete and finite. The main reason why we consider a discrete model is that there 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 $(\Delta L_t^{\mathrm{P}}, \Delta L_t^{\mathrm{T}})$.

The component of current profits that is unobservable from the point of view of the econometrician is specified as follows:

$$(6)\quad \xi_t \equiv \xi(\Delta L_t^{\mathrm{P}}, \Delta L_t^{\mathrm{T}}, \varepsilon_t) = \sigma_{\mathrm{P}}\, \varepsilon_t^{\mathrm{P}}\, \Delta L_t^{\mathrm{P}} + \sigma_{\mathrm{T}}\, \varepsilon_t^{\mathrm{T}}\, \Delta L_t^{\mathrm{T}} + \sigma_0\, \varepsilon_t^{0}(\Delta L_t^{\mathrm{P}}, \Delta L_t^{\mathrm{T}}).$$

σ_{P}, σ_{T}, and σ_0 are parameters, and we use σ_ε to represent the vector $(\sigma_{\mathrm{P}}, \sigma_{\mathrm{T}}, \sigma_0)'$. $\varepsilon_t^{\mathrm{P}}$ and $\varepsilon_t^{\mathrm{T}}$ are mutually independent standard normal random

FIGURE 6
Net Changes in Permanent Employment



Source: CBBE sample of Spanish manufacturing firms.

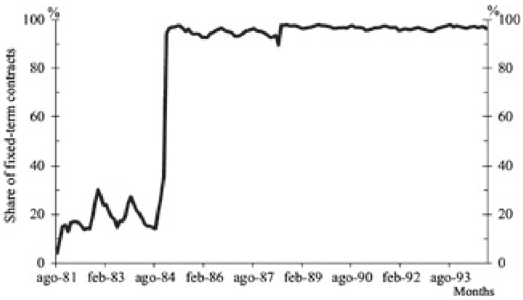
TABLE 1
Descriptive Statistics. Balanced Panel 1982–1992 (389 Firms)

	Period 1982–1984			Period 1989–1992		
Variable	**1st Quartile**	**Median**	**3rd Quartile**	**1st Quartile**	**Median**	**3rd Quartile**
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		1, 167			1, 556	

variables which are independently distributed over time and across firms. These variables try to capture unobserved heterogeneity in hiring and firing costs across firms and over time. For instance, the entitled dismissal cost of the marginal worker may vary across firms because of differences in workers' years of tenure. The variables ε_t^0 try to capture further sources of unobserved heterogeneity. 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 ε_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. Hence, the model does not hold the property of independence of irrelevant alternatives, which is a well-known limitation of the standard multinomial logit model. Nevertheless, the model retains some features of the multinomial logit, which facilitates estimation and ensures good properties of the maximum likelihood estimator.[9]

9. When $\sigma_P = \sigma_T = 0$ and $\sigma_0 > 0$, our specification becomes the one in the standard conditional logit model. When $\sigma_0 = 0$, $\sigma_P > 0$, and $\sigma_T > 0$, we have a bivariate ordered probit.

FIGURE 7
Share of Fixed-Term Contracts in Total Hiring in Spain



Source: Bover et al. (2002), Figure 1.

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 $\mathbf{x}_t \equiv (L^{\mathrm{P}}_{t-1}, L^{\mathrm{T}}_{t-1}, W_t, \eta_t)$ be the vector of state variables at period t, excluding $\varepsilon_\mathbf{t}$. And let d_t be a vector of two categorical variables representing the corresponding decisions $(\Delta L^{\mathrm{P}}_t, \Delta L^{\mathrm{T}}_t)$. The profit function can be written as:

$$\text{(7)} \qquad \Pi_t = z\,(d_t, \mathbf{x}_t)\;\theta + \xi(d_t, \varepsilon_t)$$

where θ is the 13×1 vector $(1, \theta^{\mathrm{T}}_{H0}, \theta^{\mathrm{T}}_{H1}, \theta^{\mathrm{T}}_{H2}, \theta^{\mathrm{T}}_{F0}, \theta^{\mathrm{T}}_{F1}, \theta^{\mathrm{T}}_{F2}, \theta^{\mathrm{P}}_{H0}, \theta^{\mathrm{P}}_{H1}, \theta^{\mathrm{P}}_{H2}, \theta^{\mathrm{P}}_{F0}, \theta^{\mathrm{P}}_{F1}, \theta^{\mathrm{P}}_{F2})$, and $z\,(d_t, \mathbf{x}_t)$ is a 1×13 vector with corresponding elements

(8)
$$\begin{pmatrix} Y_t - \mathrm{W}B_t,\, 1(\Delta L^{\mathrm{T}}_t > 0),\, 1(\Delta L^{\mathrm{T}}_t > 0)\Delta L^{\mathrm{T}}_t,\, 1(\Delta L^{\mathrm{T}}_t > 0)\left(\Delta L^{\mathrm{T}}_t\right)^2, \\ 1(\Delta L^{\mathrm{T}}_t < 0),\, -1(\Delta L^{\mathrm{T}}_t < 0)\Delta L^{\mathrm{T}}_t,\; 1(\Delta L^{\mathrm{T}}_t < 0)\left(\Delta L^{\mathrm{T}}_t\right)^2, \\ 1(\Delta L^{\mathrm{P}}_t > 0)\;,\; 1(\Delta L^{\mathrm{P}}_t > 0)\Delta L^{\mathrm{P}}_t\;,\; 1(\Delta L^{\mathrm{P}}_t > 0)\left(\Delta L^{\mathrm{P}}_t\right)^2\;, \\ 1(\Delta L^{\mathrm{P}}_t < 0)\;,\; -1(\Delta L^{\mathrm{P}}_t < 0)\Delta L^{\mathrm{P}}_t,\; 1(\Delta L^{\mathrm{P}}_t < 0)\left(\Delta L^{\mathrm{P}}_t\right)^2 \end{pmatrix}$$

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

$$\text{(9)} \quad V(\mathbf{x}_t, \varepsilon_t) = \max_{d_t \in D} \Bigg[z\,(d_t, \mathbf{x}_t)\,\theta + \xi(d_t, \varepsilon_t) + \beta \sum_{x_{t+1}} \int V(\mathbf{x}_{t+1}, \varepsilon_{t+1})\; f_\varepsilon(d\varepsilon_{t+1}) \times\; f_x(\mathbf{x}_{t+1}|\mathbf{x}_t, d_t) \Bigg]$$

where f_ε is the density function of $\varepsilon_\mathbf{t}$, and f_x is the transition of the vector of state variables $\mathbf{x}_t$, that has the following structure:

$$\text{(10)} \; f_x(\mathbf{x}_{t+1}|\mathbf{x}_t, d_t) = 1\left\{L^{\mathrm{P}}_t = L^{\mathrm{P}}_{t-1} + \Delta L^{\mathrm{P}}_t\right\} \times 1\left\{L^{\mathrm{T}}_t = L^{\mathrm{T}}_{t-1} + \Delta L^{\mathrm{T}}_t\right\} \times f_\eta(\eta_{t+1}|\eta_t)\; f_W(W_{t+1}|W_t)$$

The optimal solution of this dynamic labor demand model can be represented, following Aguirregabiria and Mira (2002), as the unique fixed point of a mapping in the space of conditional choice probabilities $\Pr(d_t|\mathbf{x}_t)$ (see Appendix B).

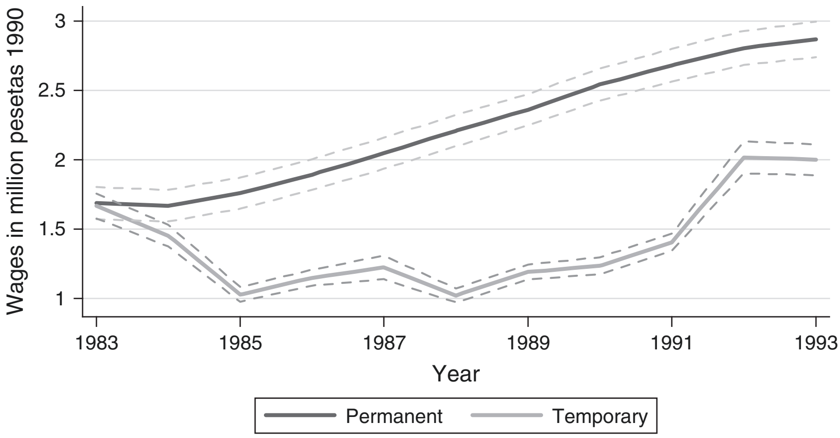
B. Assumptions for Policy Evaluation

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 is based on the estimation of our structural model and on the construction of the counterfactual steady-state distributions of employment with and without the labor market reform. Some identification assumptions are necessary to establish that we can use the pre-reform and the post-reform sample periods to construct consistent estimates of the steady-state distributions of employment with the old and the new policies. This subsection discusses these assumptions.

a) Nonanticipation 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 had 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 adjustment costs for the pre-reform period. Since our sample covers only 3 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

FIGURE 8
Time Series of the Estimated Average Wages by Contract



Note: Dotted lines denote the 95% confidence intervals.
Source: Own calculations from CBBE sample of Spanish manufacturing firms.

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 the firm learning process about the features of the new policy rule were slow. 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. This is shown in Figure 7, taken from Bover, Arellano, and Bentolila (2002), that shows how the share of fixed-term contracts in total hiring jumped from levels below 20%, just before the reform, to more than 90% just after the reform, and it remained at that level at least until 1994.

It is important to emphasize that our assumption of instantaneous learning does not imply that the new steady state was reached instantaneously after the reform. While firms' optimal decision rules on hiring and firing may have jumped to their new post-reform values soon after the implementation of the new policy, it may take several years for the distributions of employment by type of contracts to reach their new steady state. Actually, both Figures 2 and 7 are consistent with this assumption.

c) Policy-related and policy-invariant parameters. The reform entailed a change in the dismissal costs of temporary workers, θ^T_{F0}, θ^T_{F1}, θ^T_{F2}. Hiring costs of temporary workers, θ^T_{H0}, θ^T_{H1}, θ^T_{H2}, their relative productivity, λ, and hiring costs of permanent workers may have been affected by the reform as well. Therefore, θ^T_{F0}, θ^T_{F1}, θ^T_{F2}, θ^T_{H0}, θ^T_{H1}, θ^T_{H2}, θ^P_{H0}, θ^P_{H1}, θ^P_{H2} and λ are policy related parameters. It seems plausible that the technological parameter α_L, firing costs of permanent workers, θ^P_{F0}, θ^P_{F1}, θ^P_{F2}, and the stochastic process of the productivity shock are policy invariant parameters.

It must be recalled that temporary contracts were linked to the principle of causality before the reform, so that they must be aimed at jobs that were temporary in nature. However, this was no longer the case after the reform, which allowed temporary contracts for any kind of job. For this reason, we must expect that the type of workers to which temporary contracts are aimed were different before and after the reform. Our empirical model acknowledges this fact letting wages, relative productivity and firing costs to differ before and after the reform. Therefore, neither our results nor our counterfactual analysis are limited by this fact.

In an equilibrium framework, the stochastic process of wages may be affected by this reform. Our estimation of the econometric model takes this into account. However, our model is of partial equilibrium and our policy evaluation assessments provide partial equilibrium effects. We provide more discussion on this point when we present our counterfactual experiments in Section VI.

V. 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: $\{Y_{it}, L_{it}^{\mathrm{T}}, L_{it}^{\mathrm{P}}, K_{it}, I_{it}, \mathrm{W}B_{it} : i = 1, ..., N;\ t = 1, ..., T_i\}$. Our econometric model consists of 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 $\{\alpha_{\mathrm{L}}, \lambda, \theta, \sigma_{\varepsilon}, \pi, \beta\}$, where θ is (as defined in previous section) the vector of adjustment costs parameters, and π 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 $\{\alpha_{\mathrm{L}}, \lambda, \pi\}$ from the production function and transition data. In a second step, we estimate θ and σ_{ε} by maximum likelihood in the dynamic labor demand model. The maximum likelihood estimator (MLE) in the second step is a partial MLE, because the parameters in the production function have been estimated separately. Therefore, our estimation approach does not provide fully efficient estimates. However, as we show in Section V.C, our estimates of the parameters in adjustment costs are quite precise. Before describing our estimation methods and results, we undertake the estimation of wages by type of contract.

A. *Estimation of Wages*

The combination of Equations (3) and (4) implies the following regression equation in first differences:

$$\begin{aligned}(11)\qquad &\left((\mathrm{W}B_{it}/L_{it}^{\mathrm{P}}) - (\mathrm{W}B_{it-1}/L_{it-1}^{\mathrm{P}})\right)\\ &\quad = \widetilde{\gamma}_t + W_t^{\mathrm{T}}\left(L_{it}^{\mathrm{T}}/L_{it}^{\mathrm{P}}\right)\\ &\quad\quad + W_{t-1}^{\mathrm{T}}\left(L_{it-1}^{\mathrm{T}}/L_{it-1}^{\mathrm{P}}\right) + \widetilde{u}_{it}\end{aligned}$$

where $\widetilde{\gamma}_t \equiv \gamma_t - \gamma_{t-1}$, and $\widetilde{u}_{it} \equiv u_{it} - u_{it-1}$. We estimate the "parameters" $\{\widetilde{\gamma}_t, W_t^{\mathrm{T}} : t = 1983, \ldots, 1993\}$ by using OLS in Equation (11). Given the estimate $\widehat{W}_t^{\mathrm{T}}$, we can get an estimate of the wage of permanent workers in firm i as $\widehat{W}_{it}^{\mathrm{P}} = \mathrm{W}B_{it}/L_{it}^{\mathrm{P}} - \widehat{W}_t^{\mathrm{T}}\left(L_{it}^{\mathrm{T}}/L_{it}^{\mathrm{P}}\right)$. Figure 8 presents the time series of our 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 widened very importantly after 1984. This result is consistent with the evidence provided by Bentolila and Dolado (1994), by which if unions set wages for all workers and are dominated by permanent workers, then the existence of temporary contracts increases the job security and the bargaining power of permanent workers. Using longitudinal data, they find that an increase of 1 percentage point in the share of temporary contracts raises the growth rate of the wages of permanent workers by about 0.3%. Besides, Dolado, García-Serrano, and Jimeno (2002) argue that, given the weaker bargaining position of temporary workers and the higher wage pressure on permanent contracts, employers tend to under-classify temporary workers when assigning them to lower occupational categories, which determine their wage level. Also, unlike permanent workers, experience is typically not accounted in the wages of temporary workers.

B. *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. Although we maintain this assumption throughout the article, 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:

$$(12)\qquad \ln Y_{it} = \alpha_{\mathrm{K}}\ \ln K_{it} + \alpha_{\mathrm{L}}\ \ln(L_{it}^{\mathrm{P}} + \lambda\ L_{it}^{\mathrm{T}}) + \omega_{it}$$

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 of the correlation between the inputs and the unobservable productivity shock (see Griliches and Mairesse 1998). 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). Let $I_{it} = g(K_{it}, L^{\mathrm{P}}_{i,t-1}, L^{\mathrm{T}}_{i,t-1}, C_t, \omega_{it}, \chi_{it})$ be the optimal decision rule for investment in physical capital, I_{it}, where C_t represents input prices, and χ_{it} represents other possible shocks affecting investment which are unobserved to the researcher and not serially correlated and independent of the other state variables in the model. Provided that function g is strictly increasing in the productivity shock ω_{it}, there is an inverse function such that $\omega_{it} = g^{-1}(I_{it}, K_{it}, L^{\mathrm{P}}_{i,t-1}, L^{\mathrm{T}}_{i,t-1}, C_t, \chi_{it})$. Based on this expression, we can decompose ω_{it} in two additive terms: $\omega_{it} = \omega^e_{it} + \chi^*_{it}$, where $\omega^e_{it} \equiv E\left(\omega_{it} | I_{it}, K_{it}, L^{\mathrm{P}}_{i,t-1}, L^{\mathrm{T}}_{i,t-1}, C_t\right)$ 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 $\left(I_{it}, K_{it}, L^{\mathrm{P}}_{i,t-1}, L^{\mathrm{T}}_{i,t-1}, C_t\right)$, and also of L^{P}_{it} and L^{T}_{it}. Therefore, we can write the production function as,

$$(13) \quad \ln Y_{it} = \alpha_{\mathrm{L}} \ \ln(L^{\mathrm{P}}_{it} + \lambda \ L^{\mathrm{T}}_{it}) + \eta^e_{it} + \chi^*_{it}$$

where $\eta^e_{it} = \alpha_{\mathrm{K}} \ \ln K_{it} + \omega^e_{it}$. Note that η^e_{it} is a smooth function of $(I_{it}, K_{it}, L^{\mathrm{P}}_{i,t-1}, L^{\mathrm{T}}_{i,t-1}, C_t)$. 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^{P}_{it} and L^{T}_{it}, 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 $\ln Y_{it} - \widehat{\alpha}_L \ln(L^{\mathrm{P}}_{it} + \widehat{\lambda} L^{\mathrm{T}}_{it})$. According to the model, $\eta_{it} = \alpha_{\mathrm{K}} \ \ln K_{it} + \omega_{it}$. Assuming that ω_{it} follows an AR(1) process: $\omega_{it} = \rho_\omega \ \omega_{i,t-1} + a_{it}$ with $a_{it} \sim iid\,(0, \sigma^2_a)$, we have that

$$(14) \quad \left(\eta_{it} - \rho_\omega \ \eta_{i,t-1}\right) = \alpha_{\mathrm{K}} \ \left(\ln K_{it} - \rho_\omega \ \ln K_{i,t-1}\right) + a_{it}.$$

Since the innovation a_{it} is independent of $\eta_{i,t-1}$, $\ln K_{it}$ and $\ln K_{i,t-1}$, we can estimate α_{K} and ρ_ω using nonlinear least squares.

In Table 2, 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^e_{it} includes all the terms of a second order polynomial in $(I_{it}, K_{it}, L^{\mathrm{P}}_{i,t-1}, L^{\mathrm{T}}_{i,t-1})$ 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 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.

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 educated than older cohorts, and they have also accounted for a large proportion of temporary contracts.

TABLE 2

Estimation of Production Function Parameters. Unbalanced Panel 1982–1993 (2,356 Firms)[a]

Parameters	Least Squares Estimate (SE)[b]		Olley–Pakes Estimate (SE)[(2)]	
α_K	0.260	(0.006)	0.294	(0.028)
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 σ_a	0.174	(0.0025)	0.172	(0.0025)
Post-Reform σ_a	0.207	(0.0030)	0.204	(0.0030)
# Observations[c]	16,640		15,985	

[a]All the estimations include time dummies and 20 industry dummies.
[b]Standard errors are robust of heteroscedasticity and autocorrelation.
[c]In Olley–Pakes estimation we can use only those observations with investment different than zero. This explains the smaller number of observations.

The fact that the efficiency of temporary workers has increased after the reform is consistent with the increase in the wage gap between permanent and temporary workers. We are arguing that the positive effect of productivity on the relative wage of temporary workers is outweighed by a weakening in the bargaining position of temporary workers. See at the end of Section V.A above our discussion on wage bargaining and the empirical evidence in Bentolila and Dolado (1994) and Dolado, García-Serrano, and Jimeno (2002).

C. 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, have been consistently estimated in a first step, and therefore we actually observe the estimated values $\hat{W}_t^{\mathrm{T}}$, $\hat{W}_{it}^{\mathrm{P}}$, and $\hat{\eta}_{it}$. For notational convenience, we will omit the "hats." The fact that the estimated values include estimation error does not affect the consistency of our estimator of θ, though it affects its asymptotic variance.

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

(15)

$$Q(\theta, \sigma_\varepsilon, \pi, \mathrm{P}_0) = \sum_{i=1}^{N} \sum_{t=1}^{T_i} \ln \Psi(d_{it} \mid \mathbf{x}_{it};\ \theta, \sigma_\varepsilon, \pi, \mathrm{P}_0)$$

where $\Psi(d \mid \mathbf{x};\ \theta, \sigma_\varepsilon, \pi, \mathrm{P}_0)$ is a function that represents the probability that d is the optimal employment choice for the firm given that the state is $\mathbf{x}$, the structural parameters are $(\theta, \sigma_\varepsilon, \pi)$, and the firm believes that it will behave in the future according to the choice probabilities P_0. We describe the structure of the probability function Ψ in Appendix B.

Let $\hat{\mathrm{P}}_0$ be a nonparametric estimator of the set of conditional choice probabilities in P_0. And let $\hat{\pi}$ be an estimator of the parameters in the transition probability functions of wages and the productivity shock. Given these estimates, we can calculate the present values that enter into the probability function Ψ (see Appendix B) to obtain a criterion function $Q(\theta, \sigma_\varepsilon, \hat{\pi}, \hat{\mathrm{P}}_0)$ that is the log-likelihood of a random-coefficients multinomial logit model, where the random coefficients come from the term $\varepsilon^{\mathrm{P}} \Delta L^{\mathrm{P}}(d)(\sigma_{\mathrm{P}}/\sigma_0) + \varepsilon^{\mathrm{T}} \Delta L^{\mathrm{T}}(d)(\sigma_{\mathrm{T}}/\sigma_0)$. Given this likelihood, we can estimate the parameters θ, σ_0, σ_{P}, and σ_{T}. Note that these parameters are separately identified

TABLE 3
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

from θ/σ_0, σ_T/σ_0, and σ_P/σ_0 because the first element of θ, which is associated with the value of output minus the wage bill, is equal to 1. The estimator of $(\theta, \sigma_\varepsilon)$ that maximizes $Q(\theta, \sigma_\varepsilon, \hat{\pi}, \hat{P}_0)$ is consistent and asymptotically equivalent to the maximum likelihood estimator (see Proposition 4 in Aguirregabiria and Mira 2002). A recursive extension of this two-step method returns the (conditional) maximum likelihood estimator of $(\theta, \sigma_\varepsilon)$. The time-discount factor is fixed at $\beta = 0.95$.

(a) Discretization of employment changes (decision variable) and of employment levels (endogenous state variables). The main reason why we consider a discrete model of labor demand is the significant lumpiness in observed employment changes. Table 3 presents the empirical distributions of annual changes in temporary and permanent employment, i.e., in the number of workers. These variables have a discrete distribution, and a small number of discrete values account for a large proportion of observations. For instance, 57% of the observations of employment changes are between −5 and +5 workers (11 discrete values), and 72% of the observations are between −10 and +10 (21 values). We believe that a continuous-choice model of labor demand is not appropriate for this type of data. Furthermore, the solution and estimation of the dynamic decision model requires discretization of state variables.

Although the distributions of employment changes and employment levels are discrete and lumpy, they have also long tails. We would need a discrete grid with too many values to account for more than 90% of the sample values of these variables. A large dimension of the state space implies a significant increase in computational cost in the solution and estimation of the model. To deal with this issue, we consider a discretization that is firm specific. The main reason why the range of sample variation of employment variables is quite large is the existence of between-firm heterogeneity in the level of employment. Actually, the range of values in the variation of employment over time for a firm is much narrower than the range of values in the cross-sectional distribution of employment across firms. Therefore, using a firm-specific discretization, we can capture most between-firms and within-firms sample variation in employment with a relatively small and manageable state space. For $j = \mathrm{T}, \mathrm{P}$, $\Delta\tilde{L}_{it}^j$ and $\tilde{L}_{it}^j$ be the discretized values of ΔL_{it}^j and L_{it}^j, respectively. Let $\overline{L}_i$ be the "long-term" total employment of firm i as measured by the firm-specific sample mean.[10] And define the discrete variables $d_{it}^j \equiv 100(\Delta\tilde{L}_{it}^j/\overline{L}_i)$, and $\ell_{it}^j \equiv 100(\tilde{L}_{it}^j/\overline{L}_i)$. The range of values for the discretized employment changes, d_{it}^j, is defined as percentages of $\overline{L}_i$ between -20% and $+20\%$ with steps of 2 percentage points, i.e., $d_{it}^j \in \{-20, -18, \ldots, -4, -2, 0, +2, +4, \ldots, +18, +20\}$. Similarly, ℓ_{it}^j, is defined as percentages of $\overline{L}_i$ between 40% and 120%, with steps of 2 percentage points, i.e., $\ell_{it}^j \in \{40, 42, \ldots, 118, 120\}$. Given the discrete variables d_{it}^j and ℓ_{it}^j, we can obtain the values for employment changes and employment levels as $\Delta\tilde{L}_{it}^j = (d_{it}^j/100)\overline{L}_i$ and $\tilde{L}_{it}^j = (\ell_{it}^j/100)\overline{L}_i$, respectively. In Figure 9, we present the histograms of the discrete decision variables d_{it}^{P} and d_{it}^{T}, and of the endogenous state variables ℓ_{it}^{P} and ℓ_{it}^{T}.

(b) Discretization of exogenous state variables. We follow Tauchen (1986) and Tauchen and Hussey (1991) to choose the discretization grid of the exogenous state variables $(W_t^{\mathrm{T}}, W_{it}^{\mathrm{P}}, \eta_{it})$. For each of these variables, we estimate an AR(1) process and follow Tauchen-Hussey procedure. However, for the state variables $(W_{it}^{\mathrm{P}}, \eta_{it})$ 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: $W_{it} - \overline{W}_i$ and $\eta_{it} - \overline{\eta}_i$. 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 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).

To estimate our dynamic labor demand model, we have considered alternative specifications of the unobservables, including the pure conditional logit without random coefficients and different random-coefficient models, and alternative assumptions on the variances of the ε's, including homoscedasticity and heteroscedasticity. The choice of our most-preferred specification has been based on two criteria. First, the sign and magnitude of the estimated parameters should have economic sense. Second, 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, the average job turnover rates, and the cross-sectional variance of employment levels.

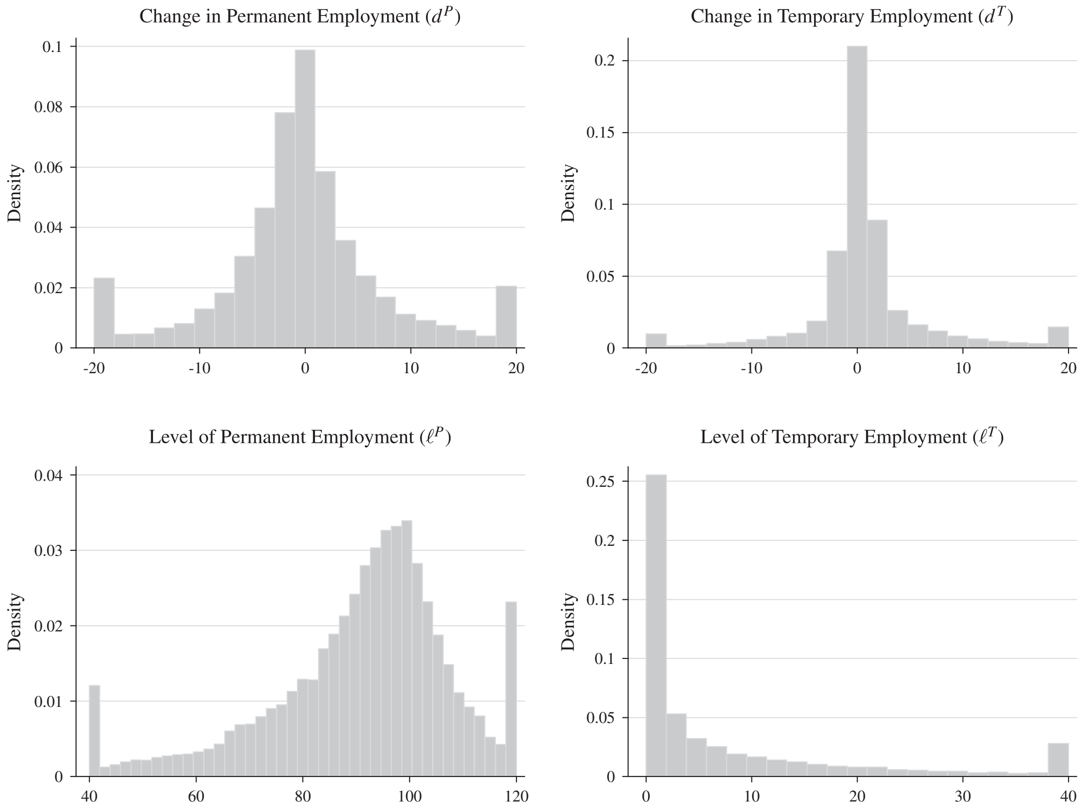
Following these criteria, our favorite specification is a model where labor adjustment costs (fixed, linear, and quadratic) and the standard deviation of the unobservable ε's are proportional to the firm-specific mean of the salary-per-worker. For instance, the linear cost of firing permanent workers for firm i is $\theta_{F1i}^{\mathrm{P}} = \tilde{\theta}_{F1}^{\mathrm{P}}\ \overline{W}_i$, where $\tilde{\theta}_{F1}^{\mathrm{P}}$ is the same parameter for every firm and $\overline{W}_i$ is firm i's mean salary per worker, i.e., $\overline{W}_i = (1/T_i)\sum_{t=1}^{T_i} W_{it}$. The same specification applies to the other θ parameters in labor adjustment costs. Similarly, the variances of the unobservables are $var(\varepsilon_{it}^{\mathrm{P}}) = \sigma_{\mathrm{P}}^2\ \overline{W}_i^2$, $var(\varepsilon_{it}^{\mathrm{T}}) = \sigma_{\mathrm{T}}^2\ \overline{W}_i^2$, and $var(\varepsilon_{it}^0) = \sigma_0^2\ \overline{W}_i^2$. It is important to note that 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. Besides, such model fails to fit the thick tails in the empirical distribution of employment changes.

Table 4 presents the estimates of the dynamic labor demand model for our preferred specification. We have estimated the model for three subperiods: the pre-reform period 1983–1984, and two post-reform periods, 1985–1988 and 1989–1992. Table 5 provides goodness-of-fit measures of the estimated model. The fit of the estimated model, for the three sample periods, to the different aggregate statistics, is very good, except for the proportion of zeros in the change of temporary employment. In this latter case, the model underestimates such proportion, yet the degree of underestimation is similar for the three sample periods.

Regarding adjustment costs, the linear components reveal as the most important ones for both contract types, either for hiring or for firing. The quadratic components are very small and nonsignificant. The fixed components are generally small, except the fixed cost of firing permanent workers in the three periods, and the fixed cost of hiring temporary workers in the pre-reform period.

10. We could have used other measure of firm size for the discretization, such as average capital stock, $\overline{K}_i$, average output, $\overline{Y}_i$, or average TFP, $\overline{\eta}_i$. All these variables are highly correlated with average employment.

FIGURE 9
Histograms of Discretized Decision and State Variables



In the case of linear hiring components, the hiring costs per worker are fairly similar for both contract types, ranging between 10% and 18% of a worker's annual salary. Interestingly, linear hiring costs for both contract types seem to experience a decline after the reform. This fact may be pointing out that the screening costs of workers are lowered after the reform. In the case of temporary workers, most new contracts become temporary contract after the reform, so that the pool of candidates for a job under a temporary contract is hugely widened. In the case of permanent workers, hiring a permanent worker after the reform means, typically, promoting a temporary worker who was already in the firm to a permanent position. Such promotion may be less expensive than the recruitment of an outsider.

The linear firing costs for permanent workers amount between 46% and 53% of a worker's annual salary, and are fairly similar between the pre-reform and the post-reform periods. Taking into account that the regulated severance payments for permanent contracts established 45 days of salary for each year of tenure, our estimate of the linear firing cost corresponds to the dismissal compensation for a worker with 4 years of tenure. This seems a reasonable tenure for the typical dismissed permanent worker. Linear firing costs for temporary workers, on the other hand, are relatively small (between 4% and 10%), showing a decline between the pre-reform and the post-reform periods, when they become nonsignificant.

Comparing the estimated adjustment cost components before and after the reform, we find several significant reductions in the post-reform period. Many of them affected the structure of adjustment cost for temporary employment: in particular, its fixed and linear components of both hiring and firing costs were reduced. The most outstanding drop corresponded to fixed hiring costs for temporary employment. This dramatic change in such costs indicates clearly the

TABLE 4

Estimation of the Dynamic Labor Demand Model. Unbalanced Panel 1982–1993 (2,356 Firms)[a]

	Period 1982–1984		Period 1985–1988		Period 1989–1992	
AC Parameters	**Permanent**	**Temporary**	**Permanent**	**Temporary**	**Permanent**	**Temporary**
Fixed hiring cost: $\theta^{\mathbf{j}}_{\mathbf{H0,i}}/\overline{W}_i$	0.012	**1.417****	0.018	0.097**	0.028	0.049
	(0.061)	**(0.060)**	(0.035)	(0.039)	(0.041)	(0.046)
Linear hiring cost: $\theta^{\mathbf{j}}_{\mathbf{H1,i}}/\overline{W}_i$	**0.183****	**0.181****	**0.117****	**0.107****	**0.101****	**0.089****
	(0.058)	**(0.049)**	**(0.038)**	**(0.041)**	**(0.043)**	**(0.045)**
Quad hiring cost: $\theta^{\mathbf{j}}_{\mathbf{H2,i}}/\overline{W}_i$	0.0003	0.0001	0.0006	−0.0001	0.0005	−0.0008
	(0.0006)	(0.0004)	(0.0008)	(0.0006)	(0.0005)	(0.0009)
Fixed firing cost: $\theta^{j}_{\mathbf{F0,i}}/\overline{W}_i$	0.083**	0.067	0.136**	0.061	0.080**	0.058
	(0.038)	(0.094)	(0.024)	(0.084)	(0.036)	(0.113)
Linear firing cost: $\theta^{j}_{\mathbf{F1,i}}/\overline{W}_i$	**0.514****	**0.098****	**0.464****	**0.060**	**0.528****	**0.051**
	(0.098)	**(0.045)**	**(0.035)**	**(0.037)**	**(0.080)**	**(0.048)**
Quad firing cost: $\theta^{\mathbf{j}}_{\mathbf{F2,i}}/\overline{W}_i$	−0.00043*	−0.00006	0.00006	0.00037	−0.0006	0.0005
	(0.00022)	(0.0011)	(0.0007)	(0.0008)	(0.0008)	(0.0009)
Other Parameters	**Period 1982–1984**		**Period 1985–1988**		**Period 1989–1992**	
$\sigma_{P,i}/\overline{W}_i$	0.611		0.515		0.540	
	(0.080)		(0.058)		(0.079)	
$\sigma_{T,i}/\overline{W}_i$	0.170		0.172		0.149	
	(0.025)		(0.019)		(0.026)	
$\sigma_{0,i}/\overline{W}_i$	0.871		0.710		0.761	
	(0.069)		(0.054)		(0.064)	
	Period 1982–1984		**Period 1985–1988**		**Period 1989–1992**	
# Observations	2, 274		7, 219		6, 257	
LR index[b]	0.232		0.220		0.267	

[a]All the parameters are unit-free because all firing and hiring costs are proportional to the firm-specific average wage $\overline{W}_i$. Standard errors are reported between parentheses.

[b]The LR (Likelihood Ratio) Index is a measure of goodness of fit 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-likelihood under the the hypothesis that all parameters except $\sigma_{0,i}$ are equal to zero.

difference in the aim of temporary contracts, and therefore in the characteristics of workers hired under such contracts, before and after the reform. Namely, the link to the causality principle (cover an absent post, seasonal work, production campaign, etc.) can justify high hiring costs before the reform. Under the limited reasons that justify the use of temporary contracts before the reform, firms may have strong incentives to screen candidates, which can explain substantial fixed costs of hiring (e.g., when hiring was aimed to cover unexpected absent posts). Together with the significant drop in the fixed cost of hiring temporary workers, we find an increase in the proportion of firms using temporary workers from 45% in 1984 to 75% in 1993, accompanied by a more intensive use by these firms of temporary contracts. Concerning permanent employment, the linear hiring and firing costs were reduced after the reform.

Finally, the parameters related with the dispersion of the unobservable shocks are all significantly positive, and keep 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; (3) the productivity of temporary workers has become closer to the one of permanents; and (4) the wage-gap between permanent and temporary workers has widened after the reform.

TABLE 5
Goodness of Fit Measures of the Estimated Model. Unbalanced Panel 1982–1993 (2,356 Firms)

Statistics	Period 1983–1984 Model (Empirical)	Period 1985–1988 Model (Empirical)	Period 1989–1992 Model (Empirical)
Median permanent employment per firm	98.0 (95.0)	64.0 (66.0)	59.0 (56.0)
Mean proportion of temporary workers	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 on $d^{\mathrm{P}} > 0$	4.0% (3.9%)	4.8% (5.2%)	5.6% (5.7%)
Median Value of d^{P} conditional on $d^{\mathrm{P}} < 0$	−3.9% (−3.8%)	−4.0% (−4.3%)	−5.1% (−5.2%)
Median Value of d^{T} conditional on $d^{\mathrm{T}} > 0$	1.7% (1.7%)	2.9% (2.7%)	4.0% (4.2%)
Median Value of d^{T} conditional on $d^{\mathrm{T}} < 0$	−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)

VI. 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 such 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 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, real 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 fix the values of the parameters at their 1983–1984 level, except for the linear firing costs θ^{P}_{F1} and θ^{T}_{F1} which are reduced by half: i.e., the counterfactual values of $\theta^{\mathrm{P}}_{F1,i}/\overline{W}_i$ and $\theta^{\mathrm{T}}_{F1,i}/\overline{W}_i$ 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 6 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 rises 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 alike (a 4.1% increase), the counterfactual reform had improved 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%).

Recently, several Spanish economists (see Andrés et al. 2009) have advocated a new labor market reform in Spain to establish a single (permanent) contract with lower firing costs. Actually, in our experiment, the proportion of

TABLE 6
Evaluation of the Labor Market Reform

Statistics	Pre-Reform Economy[a]	Post-Reform Economy[a]	Counterfactual Reform[a]
Median permanent employment per firm	99.0	89.1 (−10.0%)	105.5 (+6.6%)
Median total employment per firm	102.7	106.3 (+3.5%)	106.9 (+4.1)
Mean proportion of temporary workers	3.8%	16.2%	1.3%
Median absolute value of $\Delta\tilde{L}^{P}$	2.9%	2.5%	4.2%
Median absolute value of $\Delta\tilde{L}^{T}$	0.0%	3.1%	0.9%
Median output per firm[b]	100	100.7	101.9
Median value of a firm[b]	100	101.2	104.8

[a]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_{F1}^{P} and θ_{F1}^{T} which are reduced by half.
[b]We normalize at 100 output-per-firm and value of a firm in the pre-reform model.

temporary workers under the counterfactual is very small (1.3%). Our counterfactual policy is very similar to the advocated labor market reform if firing costs were reduced by 50%. More generally, our estimates show that a sufficiently large reduction in firing costs of permanent workers would make temporary contracts suboptimal, being equivalent to the legal elimination of this type of contract.

VII. 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. The reform had 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 halved firing costs for both contracts implies similar effects on total employment, but the positive effects on output, value of firms, and permanent employment are much stronger.

Between 1984 and 1997, the labor market regulations remained essentially unchanged. Since 1997, several reforms of labor market regulations have been undertaken. Most of them have been aimed at limiting the widespread use of temporary contracts, and boost the use of permanent contracts through lower social security and termination costs. The rules to limit temporary contracts have proved ineffective, to the extent that the proportion of temporary workers has remained stable around 30% for the total economy (with a slightly lower incidence in manufacturing). With regard to the reduction in severance payments for permanent contracts, they have been limited to new contracts of certain collectives of workers, while the severance payments for existing contracts and for many new permanent contracts have remained unchanged. The behavior of the Spanish labor market, therefore, has not lived substantial changes since the 1984 reform, and features Spain nowadays as the OECD country with the highest proportion of temporary employment.

The duality of the Spanish labor market has been strengthened in the last two decades, so that 30% of the working people, those with temporary contracts, bear most of employment turnover, to the extent that all the flexibility of the labor market is provided by them. The incidence of these contracts differs very much by worker characteristics, affecting more female, young, and less educated people, as well as people with long unemployment episodes. We have seen that the introduction of these contracts led to an increase in the employment level, but at the expense of a lower productivity per worker. In fact, unlike other countries (see Goux, Maurin, and Pauchet 2001), it seems that the role of temporary contracts as a screening device under asymmetric information is very limited. In the presence of such low conversion rates, firm-specific human capital investment seems to be negatively affected by temporary contracts. According to Dolado, García-Serrano, and Jimeno (2002), during the cyclical upturn of 1986–1990, where employment growth was mostly due to temporary contracts, labor productivity hardly reached an annual average growth of 1%. And labor productivity growth has been similarly low during the 1997–2000 upturn. The lack of firm-specific human capital investment may be behind this fact. Precisely, the counterfactual reform that we have posed illustrates this negative aspect of temporary contracts. We have seen that an alternative reform reducing all firing costs by 50% would provide a similar employment level, yet with a low proportion of temporary employment, and as a consequence, a higher productivity per worker than under the current setting. The current regulations of the Spanish labor market fail to provide incentives for firm-specific human capital investment. Several recent initiatives have advocated for the elimination of the firing cost gap between temporary and permanent contracts in Europe, introducing a single labor contract with a severance pay that would be low at the beginning of the contract and increase with workers' tenure. The last labor market reform, implemented by the Spanish Government at the beginning of 2012, has reduced dismissal costs for permanent workers who will be hired thereafter and widened the causes for fair dismissals, but use of temporary contracts has not been limited, so the dual labor market still prevails. In contrast with the existing dual schedule, the single contract proposal does not discourage job creation and firm-specific human capital investment, removes the incentives to the inefficient turnover of temporary workers and keeps protection to tenured workers.

Finally, we must acknowledge that the aforementioned data limitations may affect our estimates, and therefore our results should be qualified. First, our dataset lacks measures of gross employment flows, so that our results are based on measures net employment changes. Given the relative importance of voluntary quits and simultaneous hires and layoffs in the presence of sizeable redundancy payments, this may lead to underestimate the importance of firing costs for permanent workers. Second, our dataset covers only manufacturing firms, and overrepresents large firms. The incidence of temporary contracts is lower in manufacturing than in other sectors of the economy, and is also lower for larger firms. Regarding this, we believe that the consequences of the two-tier labor market that we have analyzed, as well as the potential effects of the counterfactual reform, are expected to be larger for the whole economy.

APPENDIX A. DATA DESCRIPTION

Construction of the CBBE Dataset

The sample consists of an unbalanced panel of nonenergy manufacturing firms, recorded in the database of the Bank of Spain's Central Balance Sheet Office. This dataset was started in 1982 collecting firm data on a voluntary basis, which implied that largest firms were overrepresented in the sample. However, the tendency in subsequent years has been characterized by the addition of firms of smaller relative size. The firms included in this database represent almost 40% of the total value added in Spanish manufacturing, so it accounts for a very large proportion of output in the Spanish economy. In Table A1, we present the sample distribution of firms by industry and size (measured as the time average of firm's employees).

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. Also, we have disregarded those firms with a public share above 50%. Finally, for the purposes of estimation, we have required firms to have at least 5 years of consecutive observations. Under these criteria, we have discarded 12% of the observations. In terms of total number of employees, the average employment is very similar in the final sample and in the raw data (248 vs. 238 employees, respectively).

Construction of Variables

Employment. Employees are disaggregated by type of contract in permanent (open-ended) and temporary (fixed-term) employees. Information on the average duration (in weeks) of temporary contracts is also available. To maintain measurement consistency, the number of temporary employees is normalized 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 happens in most firm-level datasets, there is not information on employment flows along the year, and therefore only net employment changes are observed.

Wages. 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.

Output. 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-digit industry level.

TABLE A1
Distribution of Firms by 2-digit Industry and by Size. Unbalanced Panel 1982–1993 (2,356 Firms)

		Small	Med1	Med 2	Large	Total
Iron, steel, and metal (22)	Abs. freq.	5	8	10	22	45
	% by ind.	11.11	17.78	22.22	48.89	100.00
	% by size	1.29	0.94	1.73	4.10	1.91
Bldg. materials, glass, and ceramics (24)	Abs. freq.	27	88	34	33	182
	% by ind.	14.84	48.35	18.68	18.13	100.00
	% by size	6.98	10.29	5.89	6.15	7.72
Chemicals (25)	Abs. freq.	39	99	76	92	306
	% by ind.	12.75	32.35	24.84	32.07	100.00
	% by size	10.08	11.58	13.17	17.13	12.99
Nonferrous metal (31)	Abs. freq.	38	103	53	31	225
	% by ind.	16.89	45.78	23.56	13.78	100.00
	% by size	9.82	12.05	9.19	5.77	9.55
Basic machinery (32)	Abs. freq.	29	52	47	33	161
	% by ind.	18.01	32.30	29.19	20.50	100.00
	% by size	7.49	6.08	8.15	6.15	6.83
Office machinery (33)	Abs. freq.	0	1	0	3	4
	% by ind.	0.00	25.00	0.00	75.00	100.00
	% by size	0.00	0.12	0.00	0.56	0.17
Electric materials (34)	Abs. freq.	11	29	24	35	99
	% by ind.	11.11	29.29	24.24	35.35	100.00
	% by size	2.84	3.39	4.16	6.52	4.20
Electronic (35)	Abs. freq.	3	8	10	14	35
	% by ind.	8.57	22.86	28.57	40.00	100.00
	% by size	0.78	0.94	1.73	2.61	1.49
Motor vehicles (36)	Abs. freq.	8	21	25	36	90
	% by ind.	8.89	23.33	27.78	40.00	100.00
	% by size	2.07	2.46	4.33	6.70	3.82
Ship building (37)	Abs. freq.	3	2	2	6	13
	% by ind.	23.08	15.38	15.38	46.15	100.00
	% by size	0.78	0.23	0.35	1.12	0.55
Other motor vehicles (38)	Abs. freq.	2	5	5	6	18
	% by ind.	11.11	27.78	27.78	33.33	100.00
	% by size	0.52	0.58	0.87	1.12	0.76
Precision instruments (39)	Abs. freq.	2	8	3	4	17
	% by ind.	11.76	47.06	17.65	23.53	100.00
	% by size	0.52	0.94	0.52	0.74	0.72
Nonelaborated food (41)	Abs. freq.	53	83	46	48	230
	% by ind.	23.04	36.09	20.00	20.87	100.00
	% by size	13.70	9.71	7.97	8.94	9.76
Food, tobacco, and drinks (42)	Abs. freq.	53	51	31	45	180
	% by ind.	29.44	28.33	17.22	25.00	100.00
	% by size	13.70	5.96	5.37	8.38	7.64
Basic textile (43)	Abs. freq.	20	57	53	37	167
	% by ind.	11.98	34.13	31.74	22.16	100.00
	% by size	5.17	6.67	9.19	6.89	7.09
Leather (44)	Abs. freq.	4	16	12	4	36
	% by ind.	11.11	44.44	33.33	11.11	100.00
	% by size	1.03	1.87	2.08	0.74	1.53
Garment (45)	Abs. freq.	11	48	34	22	115
	% by ind.	9.57	41.74	29.57	19.13	100.00
	% by size	2.84	5.61	5.89	4.10	4.88
Wood and furniture (46)	Abs. freq.	21	45	26	8	100
	% by ind.	21.00	45.00	26.00	8.00	100.00
	% by size	5.43	5.26	4.51	1.49	4.24

TABLE A1
Continued

		Small	Med1	Med 2	Large	Total
Cellulose and paper edition (47)	Abs. freq.	29	63	42	33	167
	% by ind.	17.37	37.72	25.15	19.76	100.00
	% by size	7.49	7.37	7.28	6.15	7.09
Plastic materials (48)	Abs. freq.	22	46	33	17	118
	% by ind.	18.64	38.98	27.97	14.41	100.00
	% by size	5.68	5.38	5.72	3.17	5.01
Other nonbasic (49)	Abs. freq.	7	22	11	8	48
	% by ind.	14.58	45.83	22.92	16.67	100.00
	% by size	1.81	2.57	1.91	1.49	2.04
Total	Abs. freq.	387	855	577	537	2,356
	% 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: Size is defined in accordance with the firm's average number of employees, which is up to 25 for Small, between 25 and 75 for Med1, between 75 and 200 for Med2, and greater than 200 for Large.

Physical capital. We are interested in depreciable physical capital which is already productive, so Land and Capital stock in course of construction are excluded from the definition of the stock of physical capital. Since the CBBE does not have independent estimates of investment available, gross nominal investment I_{it} must be imputed from changes in the book value of physical capital with a correction for depreciation, that is $I_{it} = KNB_{it} - KNB_{i,t-1} + \text{Dep}_{it} + Rev_{it}$ where, $KNB_{it} = KGB_{it} - A\text{Dep}_{it}$ is the book value of the net stock of physical (book value of the gross stock of physical capital KGB_{it} minus accumulated depreciation $A\text{Dep}_{it}$); Dep_{it} is the accounting depreciation during the year; and Rev_{it} is the net variation in the book value of physical capital and in its accumulated depreciation due to positive and/or negative revaluations. To calculate the replacement value of capital, we use a perpetual inventory method which takes account for depreciation and inflation. To do this, an initial value for the first year that data is available for a given firm is calculated as $q_1 K_{i1} = (q_1/q_{1-AA_i}) \times KGB_{it}(1-\delta_i)^{AA_i}$ where q_t is the price deflator of the stock of physical capital at year t; δ_i is the average depreciation rate of the stock of physical capital; and AA_i is the average age of the stock of physical capital, which is approximated by the ratio $\text{ADep}_{i1}/\text{Dep}_{i1}$ for the first year in which data for the firm are available. Furthermore, the average depreciation rate is computed at the firm level as the ratio of firm's average accounting depreciation to the firm's average accumulated depreciation. As regards price indices, the corresponding GDP implicit deflator of investment goods is used (Source: INE). The recursive method to compute the replacement value of the stock of physical capital from the second year that data is available is $q_t K_{it} = (q_t/q_{t-1}) \times K_{i,t-1}(1-\delta_i) + I_{it}$,which assumes that investment occurs at the end of the year.

APPENDIX B. PROBABILISTIC REPRESENTATION OF FIRMS' EMPLOYMENT DECISIONS

Remember that the specification of the one-profit function is $\Pi_t = z(d_t, \mathbf{x}_t)\ \theta + \xi(d_t, \varepsilon_t)$, where θ is the 13×1 vector $\left(1, \theta^{\text{T}}_{H0}, \theta^{\text{T}}_{H1}, \theta^{\text{T}}_{H2}, \theta^{\text{T}}_{F0}, \theta^{\text{T}}_{F1}, \theta^{\text{T}}_{F2}, \theta^{\text{P}}_{H0}, \theta^{\text{P}}_{H1}, \theta^{\text{P}}_{H2}, \theta^{\text{P}}_{F0}, \theta^{\text{P}}_{F1}, \theta^{\text{P}}_{F2}\right)$, and $z(d_t, \mathbf{x}_t)$ is a 1×13 vector as described in Equation (8). The unobservable term $\xi(d_t, \varepsilon_t)$ is:

$$\text{(A1)}\quad \xi(d_t, \varepsilon_t) = \sigma_{\text{P}}\ \varepsilon^{\text{P}}_t\ \Delta L^{\text{P}}(d_t) + \sigma_{\text{T}}\ \varepsilon^{\text{T}}_t \Delta L^{\text{P}}(d_t) + \sigma_0\ \varepsilon^0_t(\Delta L^{\text{P}}(d_t), \Delta L^{\text{T}}(d_t))$$

where $\Delta L^{\text{P}}(d)$ and $\Delta L^{\text{T}}\text{T}(d)$ represents the values of ΔL^{P}_t and ΔL^{T}_t associated to the discrete choice alternative $d_t = d$. If we had a "myopic" model (i.e., $\beta = 0$), the choice probabilities in the likelihood function would have the following form:

$$\text{(A2)}\quad \Pr(d \mid \mathbf{x}) = \int \frac{\exp\left\{z(d, \mathbf{x})\,\frac{\theta}{\sigma_0} + \varepsilon^{\text{P}} \Delta L^{\text{P}}(d)\frac{\sigma_{\text{P}}}{\sigma_0} + \varepsilon^{\text{T}} \Delta L^{\text{T}}(d)\frac{\sigma_{\text{T}}}{\sigma_0}\right\}}{\sum_{j \in D} \exp\left\{z(j, \mathbf{x})\,\frac{\theta}{\sigma_0} + \varepsilon^{\text{P}} \Delta L^{\text{P}}(j)\frac{\sigma_{\text{P}}}{\sigma_0} + \varepsilon^{\text{T}} \Delta L^{\text{T}}(j)\frac{\sigma_{\text{T}}}{\sigma_0}\right\}} \times \phi(d\varepsilon^{\text{P}})\phi(d\varepsilon^{\text{T}})$$

where ϕ is the pdf of the standard normal. We can calculate numerically the double integral in this probability function. In this article, we have calculated this integral using the Gauss-Legendre quadrature method provided by the command *intquad2* in the GAUSS software package.

In our dynamic, forward-looking, model, the choice probabilities have a similar structure as in (A2) but we have to replace the current profit $z(d, \mathbf{x})\,\theta/\sigma_0$ by the present value (expected and discounted intertemporal profits) of choosing alternative d when the state is $\mathbf{x}$. 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* $\text{P} = \{P(d|\mathbf{x}) : (d, \mathbf{x}) \in D \times X\}$, where $P(d|\mathbf{x})$ is the probability that the optimal decision is d conditional on the value of $\mathbf{x}_t$ being $\mathbf{x}$. Define also π as the vector of parameters that characterize the transition probabilities f_η and f_W. We can obtain the vector P as the unique fixed point of a contraction mapping in the space of conditional choice probabilities: $\text{P} = \Psi(\text{P})$. Under our assumptions on the probability distribution of ε_t, this mapping is:

(A3)

$$\Psi(d \mid \mathbf{x}; \mathrm{P}) = \int \frac{\exp\left\{\begin{array}{l} Z_{\mathrm{P}}(d,\mathbf{x})\dfrac{\theta}{\sigma_0} + e_{\mathrm{P}}(d,\mathbf{x}) \\ +\varepsilon^{\mathrm{P}}\Delta L^{\mathrm{P}}(d)\dfrac{\sigma_{\mathrm{P}}}{\sigma_0} + \varepsilon^{\mathrm{T}}\Delta L^{\mathrm{T}}(d)\dfrac{\sigma_{\mathrm{T}}}{\sigma_0} \end{array}\right\}}{\sum_{j\in D}\exp\left\{\begin{array}{l} Z_{\mathrm{P}}(j,\mathbf{x})\dfrac{\theta}{\sigma_0} + e_{\mathrm{P}}(j,\mathbf{x}) \\ +\varepsilon^{\mathrm{P}}\Delta L^{\mathrm{P}}(j)\dfrac{\sigma_{\mathrm{P}}}{\sigma_0} + \varepsilon^{\mathrm{T}}\Delta L^{\mathrm{T}}(j)\dfrac{\sigma_{\mathrm{T}}}{\sigma_0} \end{array}\right\}} \times \phi(d\varepsilon^{\mathrm{P}})\phi(d\varepsilon^{\mathrm{T}}).$$

$Z_{\mathrm{P}}(d,\mathbf{x})\dfrac{\theta}{\sigma_0} + e_{\mathrm{P}}(d,\mathbf{x})$ is the present value of choosing alternative d when the state is $\mathbf{x}$. More specifically, $Z_{\mathrm{P}}(d,\mathbf{x})$ is a 1×13 vector with the present values of the stream of current and future values of the vector $z(.,.)$ 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 P. For instance, the first element of the vector $Z_{\mathrm{P}}(d,\mathbf{x})$ is the present value of output minus wage bill. In a similar manner, $e_{\mathrm{P}}(d,\mathbf{x})$ is a scalar containing the present value of the stream of future realizations of $\xi(d_{t+j}, \varepsilon_{t+j})$ associated with optimal future choices. More formally, we have that:

$$\text{(A4)}\quad Z_{\mathrm{P}}(d,\mathbf{x}) = z(d,\mathbf{x}) + \sum\nolimits_{s=1}^{\infty} \beta^j\ \mathbb{E}\big(z\big(d^*_{t+s}, \mathbf{x}_{t+s}\big) \mid d_t = d, \mathbf{x}_t = \mathbf{x};\ \mathrm{P}\big)$$

and

$$\text{(A5)}\quad e_{\mathrm{P}}(d,\mathbf{x}) = \sum\nolimits_{s=1}^{\infty} \beta^j \times \mathbb{E}\left(\xi(d^*_{t+s}, \varepsilon_{t+s}) \mid d_t = d, \mathbf{x}_t = \mathbf{x};\ \mathrm{P}\right)$$

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

In what follows, we outline the procedure to calculate the values of $Z_{\mathrm{P}}(d,\mathbf{x})$ and $e_{\mathrm{P}}(d,\mathbf{x})$, which is explained in depth by Aguirregabiria and Mira (2002, 2010). Let $W_z^{\mathrm{P}}(\mathbf{x})\theta + W_e^{\mathrm{P}}(\mathbf{x})$ be the expected discounted utility of behaving according to choice probabilities 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 $\mathbf{W}_z^{\mathbf{P}} \equiv \{W_z^{\mathrm{P}}(\mathbf{x}) : \mathbf{x} \in X\}$ and the vector $\mathbf{W}_e^{\mathrm{P}} \equiv \{W_e^{\mathrm{P}}(\mathbf{x}) : x \in X\}$. It can be shown that $\mathbf{W}_z^{\mathrm{P}}$ is the unique solution of the recursive equation $\mathbf{W}_z^{\mathrm{P}} = \sum_{d\in D}\mathrm{P}(d) * \{\mathbf{z}(d) + \beta\ \mathbf{F}_x(d)\ \mathbf{W}_z^{\mathrm{P}}\}$, where $\mathrm{P}(d)$ 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. Likewise, $\mathbf{W}_e^{\mathrm{P}}$ is the unique solution of the recursive equation $\mathbf{W}_e^{\mathrm{P}} = \sum_{d\in D}\mathrm{P}(d) * \{\mathbf{e}^{\mathrm{P}}(d) + \beta\ \mathbf{F}_x(d)\ \mathbf{W}_z^{\mathrm{P}}\}$ where $\mathbf{e}^{\mathrm{P}}(d)$ is the vector $\{e^{\mathrm{P}}(d,\mathbf{x}) : \mathbf{x} \in X\}$ and $e^{\mathrm{P}}(d,\mathbf{x}) \equiv \mathbb{E}(\xi(d_t^*, \varepsilon_t)|d_t^* = d, \mathbf{x}_t = \mathbf{x})$, which is a known function of the choice probabilities $P(d|\mathbf{x})$. The objects $\mathbf{W}_z^{\mathrm{P}}$ and $\mathbf{W}_e^{\mathrm{P}}$ can be computed by successive approximations, iterating on the contraction mappings that implicitly define them. Note that the mappings that implicitly define $\mathbf{W}_z^{\mathrm{P}}$ and $\mathbf{W}_e^{\mathrm{P}}$ are linear in these objects. Therefore, there is a closed form expression for $\mathbf{W}_z^{\mathrm{P}}$ and for $\mathbf{W}_e^{\mathrm{P}}$. For instance, $\mathbf{W}_z^{\mathrm{P}} = (\mathbf{I} - \beta\sum_{a=0}^{J}\mathrm{P}(a) * \mathbf{F}_x(a))^{-1}\sum_{a=0}^{J}\mathrm{P}(a) * \mathbf{z}(a)$. When the number of cells in X is small enough, matrix inversion algorithms may be preferable to successive approximations. The matrix $(\mathbf{I} - \beta\mathbf{F})^{-1}$ can also be approximated using the series $\mathbf{I} + \beta\mathbf{F} + \beta^2\mathbf{F}^2 + \cdots + \beta^K\mathbf{F}^K$, with K large enough. This can be easier than matrix inversion. More generally, this inverse matrix can be obtained iterating in $\mathbf{A}$ (successive approximations) in the mapping $\mathbf{A} = \mathbf{I} + \beta\mathbf{F}\ \mathbf{A}$.

Given our definition of the state variables in deviations with respect to firm-specific means, the matrix $\mathbf{W}_z^{\mathrm{P}}$ and the vector $\mathbf{W}_e^{\mathrm{P}}$ 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 (one for each firm), each with a dimension of 6, 888 discrete values in the state space. For the calculation of the values $e^{\mathrm{P}}(d,\mathbf{x})$ 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 $e^{\mathrm{P}}(d,\mathbf{x})$ is the one that comes from the expectation of the future extreme value error ε^0.

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