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Abstract

We estimate models of labour demand for a panel of 3,400 Spanish manufacturing firms over the period 1985-2001. We examine the roles of flexible labour through temporary contracts, financial factors and a policy reform in 1997 affecting permanent contracts by lowering payroll taxes and dismissal costs. Compared to permanent employment, the demand for flexible labour displays (i) greater sensitivity to financial factors (ii) greater cyclical sensitivity (iii) a larger average wage elasticity (iv) less inertia. Our analysis of the 1997 policy reform finds an effect of payroll taxes on employment. A 5 percentage point reduction in the payroll tax increases labour demand by 8 per cent.

JEL Codes: J23, J32.

Key words: labour demand; financial pressure; temporary contracts.

1 Introduction

Labour flexibility remains a fiercely debated issue. It is commonly alleged that Europe's lack of labour flexibility has hindered its economic growth and labour market performance resulting in excessive and persistent unemployment. This is most often attributed to what are deemed excessive costs incurred in shedding labour and high payroll taxes.[1] Within Europe, the case of Spain has merited closest attention. This partly reflects the view that Spain has been the most acute example of the European unemployment problem, associated with its traditionally high levels of employment protection and particularly high unemployment rates.[2] A second important reason for wider interest in the Spanish experience stems from the series of reforms that have been introduced. These reforms have influenced policy elsewhere in Europe, notably in France and Italy.

Reform in Spain was initiated through the liberalisation of flexible contracts. From the mid-1980s, temporary contracts in Spain experienced their most widespread adoption of any industrialised economy, with almost one-third of employees employed under a temporary contract during the 1990s.[3] However, during this latter period, an era of continued high unemployment, it has been argued that a dual labour market may have developed based upon the possession of a permanent versus temporary labour contract with conversion rates from the latter to the former being low. This may in turn have had perverse effects on wage determination, impeding any flexibility benefits of

[1] For a review of this debate see OECD (1999) and Nickell and Layard (1999).

[2] In fact, a majority of European countries have unemployment rates lower than that of the United States at the time of writing. The high regulation and unemployment story at best only fits the largest four European economies of Germany, France, Italy and Spain (Nickell, 2002).

[3] Temporary work agencies have been permitted since 1994.

temporary contracts being translated into improved employment performance. Partly in response to this, policy has since reduced the appeal to employers of opting for temporary contracts (Dolado *et al.* 2002). To some extent this was achieved through a policy reform of 1997 which reduced payroll taxes and dismissal costs associated with permanent contracts.

Motivated by such issues, this paper examines labour demand at the firm-level, focusing on the case of Spain and the distinction between temporary and permanent labour. The starting point to the paper is an analysis by Bentolila and Saint-Paul (1992), who considered the consequences of temporary contracts for labour demand, also focusing on the experience of Spain. Their empirical analysis was conducted for a panel of firms but restricted to the sample period 1985-88. Our first contribution is to extend the sample period examined for a similar panel of manufacturing firms but to cover the period 1985-2001. This extension is not inconsequential since a central hypothesis highlighted by Bentolila and Saint-Paul (1992) is the different behaviour of demand for the two labour contract types over the economic cycle. To include a period of cyclical downturn, as was the case in Spain in 1993, is therefore a useful extension.

Second, and more significantly, we consider the effects of the labour reform of 1997. Dolado *et al.* (2002) suggest that the 1997 reform may have been important in reducing the unemployment rate in Spain during the late-1990s. This reform has as yet received much less attention to date than the earlier reform of 1984 which was responsible for the major liberalisation of flexible contracts. Moreover in its scope and anticipated effects the latter reform was rather different. Rather than producing an alternative to offering employment under a permanent contract, as under the 1984 reform, the reform of 1997 focused on altering the conditions of those (subsequently) hired under a permanent contract, specifically by reducing payroll taxes and dismissal costs. The difference in approach reflected a desire to reduce the insulation of permanent contract staff, partly provided by the existence of a buffer of temporary workers, to wage settlements.

We focus on the effects of the payroll tax reduction for labour demand, which constituted an important part of the 1997 reform in Spain. To date, a consensus has yet to emerge on the consequences of payroll taxes (and labour taxes more generally) for labour demand and employment. Previous studies have been largely based on aggregate evidence drawn from international comparisons. Daveri and Tabellini (2000) for instance, argue that much of the increase in unemployment across industrialized countries, particularly in continental Europe between 1965 and 1995, can be explained by increasing labour taxes. Other macroeconomic evidence has found modest effects from labour taxes to unemployment (see Nickell and Layard, 1999). In this paper, we provide microeconomic evidence on the effects of payroll taxes on labour demand exploiting variation across firms in the effects of the 1997 reform on Spanish firms.

Third, the potential role of financial factors in influencing labour demand is highlighted. Despite the large number of studies of labour demand, very few studies consider a role for financial variables. Notable exceptions are Nickell and Wadhwani (1991) and Nickell and Nicolitsas (1999), both studies of UK quoted firms, and Ogawa (2003), a study of Japanese firms. This limited number of studies that consider a role for financial variables might seem surprising given the large number of studies of fixed investment which do find a role for financial variables (see Hubbard, 1998). An additional aim of our analysis is to consider how financial factors differentially affect the demand for permanent and temporary employment. To what extent is the kind of employment response to financial pressure highlighted by Nickell and Nicolitsas (1999), borne by those on temporary contracts in an economy where such contracts have been widely adopted?

Fourth, we therefore examine the demand for flexible labour explicitly. This allows the consideration of a number of additional hypotheses from the Bentolila and Saint-Paul (1992) model that they did not consider or for which they were unable to uncover empirical support given the short time dimension of their panel of firms. This includes the predictions that temporary employment should be less persistent than that of permanent staff, and that the responsiveness to wage costs should be greater. In so doing we aim to shed further light on the consequences of flexible contracts for the labour market.

The remainder of the paper is organised as follows. Section 2 outlines the Spanish labour market including the reform of 1997 and, drawing on the theoretical model of Bentolila and Saint-Paul (1992), describes the hypotheses of interest. Section 3 presents data description and estimation results for our panel of 3,400 Spanish manufacturing firms for the period 1985-2001. Section 4 concludes.

2 Economic background

This section first describes the general characteristics of the Spanish labour market and relevant policy reforms undertaken in Spain. In arriving at a set of hypotheses we wish to confront with our company-level data, we also summarise the implications of a model presented by Bentolila and Saint-Paul (1992) which considers the consequences for labour demand of flexible contracts.

Unemployment in Spain has been amongst the highest in Europe since the early 1980s, increasing from 7 per cent in 1978 to 21 per cent by 1985, with a particularly high incidence of long-term unemployment. Recognition of the poor performance of the Spanish labour market led to a number of reforms. Moreover the Spanish approach to reform has since been adopted by other European countries. France and Italy have since pursued a similar approach of attempting to increase labor market flexibility through the liberalisation of temporary contracts.

The highly regulated nature of the Spanish labour market reflects above all the degree of employment protection. Since a number of descriptions of the institutional background in Spain exist (see for example, Jimeno and Toharia, 1994; Dolado *et al.* 2002), we offer only a broad outline here. Many of the institutional characteristics of the Spanish labour market at least through the 1980s, were inherited from the Franco regime. Although the transition to democracy from 1978 involved the legalisation of trade unions and the introduction of a relatively generous unemployment benefit system, the rules governing employment protection remained largely intact. In 1984, a major reform was undertaken in an attempt to enhance flexibility through the substantial liberalisation of temporary contracts, previously restricted to seasonal work. Initially, temporary contracts could be signed for up to 6 months at a time and renewed for up to 3 years, with low or zero dismissal costs, whilst their non-renewal or conversion to permanent status could not be appealed to tribunals. In contrast, the process of appeal to labour courts under permanent contracts frequently imposed substantial costs on employers. Under such contracts firing costs have depended upon the seniority of the employee and on the reason for dismissal. Those with such contracts had a right to appeal and this appeal procedure has played an important role in determining actual dismissal costs.[4] For instance, from 1994 dismissals for economic or technological reasons were permitted but even in this case if the dismissal was judged as unfair then the worker was entitled to reinstatement or (for contracts signed before 1997) 45 days' salary per year of tenure up to a maximum of 42 months' salary as well as the foregone salary during the period of job loss. Even if it was judged fair then the worker received a severance payment of 20 days wages' per year of tenure up to a maximum of 12 months' salary. The probability of an unfair verdict is estimated as 0.72 by Galdón-Sánchez and Güell (2000).[5]

It was subsequently realised that the two-tier labour market encouraged by the introduction and widespread adoption of temporary contracts–which by 1990 represented around one-third of employees–had likely resulted in a number of distortions. In particular it was argued that temporary workers were used as a buffer against adverse shocks and that the interests of permanent workers dominated wage negotiations and

[4] Malo (2000) emphasises that actual dismissal costs incurred by the firm exceed the 'fair' statutory minimum. For instance, a majority of workers dismissed for disciplinary or general economic reasons have sued their employer and an out of court settlement has typically followed (see Malo (2000) for an analysis). Average actual severance payments increased for almost 5 months salary in 1981 to over 12 months' salary by 1993 (Dolado and Jimeno, 1997).

[5] The system of wage negotiation, being highly centralised, reinforces the lack of flexibility (see Bover *et al.* 2002).

these workers ('insiders') were excessively insulated from the employment consequences of wage settlements (eg. Bentolila and Dolado, 1994; Dolado and Jimeno, 1997). Reforms introduced since the mid-1990s have generally aimed to reduce the attractiveness of offering employment on a temporary contract. The most significant of these reforms took place in 1997.

The 1997 reform reduced payroll taxes by between 40 and 90 per cent for (new) permanent contracts and for conversions of temporary into permanent contracts for large groups of workers, notably those aged under 30 and over 45, the long-term unemployed, and the disabled.[6] It also reduced dismissal costs for unfair dismissals for economic reasons by around 25 per cent (ie. to 33 days' salary per year of tenure). Kugler *et al.* (2002) estimate that the reform reduced labour costs (of young workers) in the first year of the permanent contract by 9.9 per cent and 9.5 per cent for men and women, respectively and in the second year by 7.9 per cent and 7.3 per cent respectively. Around 80 to 90 per cent of the reduction in labour costs is attributed to the payroll tax reduction and it is this aspect of the overall reform that we focus on. As emphasised by Kugler *et al.* (2002), the 1997 reform was by no means a marginal one.

To date, much of the evidence on flexible/temporary contracts has focused upon the experience of Spain. Most of these studies have been presented in the light of the 1984 reform. Kugler *et al.* (2002) examine the impact of the 1997 reform in terms of employment and worker flows from Spanish labour force survey data. They find evidence that the reform increased permanent employment probabilities for younger employees.

[6] The reduction in payroll taxes lasted for the first two years of a new permanent contract and one year for a conversion to a permanent contract.

Results for other types of workers finds insignificant effects. Our analysis instead adopts a firm-level approach, looking more explicitly at its effects on the demand-side of the labour market.

As a framework for understanding labour demand in a labour market with permanent and flexible contracts, Bentolila and Saint-Paul (1992) present a model in which there is labour heterogeneity in the form of two types of labour, rigid labour (Type I) and flexible labour (Type II). Workers of the two types are remunerated at rates of w_1 and w_2 respectively. In the event of redundancy, Type I workers will impose a cost f on the employer, whilst flexible labour can be dismissed without incurring such an adjustment cost. It is assumed that flexible workers are less productive perhaps because the firm itself has a stronger incentive to invest in firm-specific skills where the worker is on a permanent contract. A flexible worker is equivalent in productivity terms to ρ rigid employees ($\rho < 1$). Those on permanent contracts however, are thought to be able to exploit this greater insider power and the existence of f turnover costs when replacing them, by bargaining higher wages ($w_1 > w_2$). This wage is however, taken to be predetermined and not considered explicitly as part of the model.

Given availability of flexible contracts, the preferred solution requires that rigid workers are preferable to flexible workers where the firing cost is not incurred (ie. $w_1 < w_2/\rho$) and that flexible workers are preferred to rigid workers during an expansion where the firm may have to incur the turnover cost f should a downturn materialise (ie. $w_2/\rho < w_1 + \gamma f$). The solution obtained involves only flexible workers being recruited during expansions since the marginal cost is less than that expected under recruitment of employees on rigid contracts. Moreover, in a recession, no temporary contract workers will be employed, nor will rigid workers be made redundant during a downturn. Each firm is subject to idiosyncratic and aggregate demand shocks which govern the likelihood of entering a downturn. The burden of adjustment to cyclical effects therefore falls on temporary contract workers.

The central implications which emerge from the model are the following.

- Flexible labour is more responsive to company-specific demand shocks than is the demand for rigid/permanent labour.
- The response of rigid employment to an aggregate recession is smoother than the response of flexible employment.
- A lower degree of persistence for temporary employment, following from the lower adjustment costs incurred through adjusting temporary employment.
- The wage elasticity of labour demand is (in absolute terms) greater in the case of flexible than rigid labour. In the available data, the closest we can get to testing this hypothesis, is to consider the responsiveness of the two types of labour to the *average* labour cost at the firm.[7]

Bentolila and Saint-Paul's (1992) empirical analysis consisted of employing a firm-level panel, from the same source that we consider below, but for a much more restricted time dimension that for 1985 to 1988. Two central predictions are that labour demand should be more sensitive to demand shocks after the policy reform of 1984

[7] This also precludes estimating the substitution of one labour type for another.

and that employment should display less persistence. Bentolila and Saint-Paul (1992) consider these by interacting the demand shock variable (sales growth) and lagged employment term with a dummy for the 1986-88 period on the grounds that temporary contracts were much more widely adopted in this period. The evidence supported the proposition of a greater sensitivity to demand but did not suggest any reduction in persistence. A number of other hypotheses were also considered. Since temporary employment is expected to represent a larger share of total employment when a company is expanding, they considered variation in the labour demand equation over the (company-specific) cycle. There was evidence that employment is less persistent for companies with higher rates of growth but they find no evidence that labour demand of companies experiencing higher rates of growth is any more sensitive to the demand shock (the opposite is found). It is suggested that this may be because the demand variable also picks up a role for liquidity or financial constraints. Since financial constraints were not investigated directly little more could be said on this hypothesis. Note that a more direct way to consider a number of these hypotheses is to estimate the demand for temporary labour directly, which Bentolila and Saint-Paul (1992) do not do, instead focusing on variations between companies growing at different rates.

In addition to the above factors, and going beyond the considerations of the Bentolila and Saint-Paul (1992) analysis, we also wish to investigate the possible role of financial factors in influencing labour demand. In this context the particular hypothesis we wish to consider is the following.

- If financial factors also influence labour demand (as in Nickell and Wadhwani, 1991), again the burden of adjustment is expected to fall disproportionately on flexible labour implying a greater responsiveness to financial factors.

Our discussion of the 1997 policy reform motivates the following.

- In order to address the effects of the payroll tax reduction, we construct a measure of the average payroll tax paid by each firm and how this was affected by the 1997 reform. The reduction in the payroll tax for new contracts affecting certain groups of workers implies that this will have varied across firms and we consider this in our models for labour demand.

In the next section the analysis considers in turn each of the hypotheses described.

3 Data and Estimation

The paper confronts the hypotheses described above with data from a sample of 3,400 manufacturing firms in Spain. The labour demand equation, derived by Nickell and Nicolitsas (1999) from a quadratic adjustment cost model which then adds financial factors, takes the following form.

$$n_{it} = \alpha_i + \beta_1 n_{it-1} + \beta_2 n_{it-2} + \beta_3 w_{it-1} + \beta_4 \Delta w_{it} + \beta_5 k_{it} + \beta_6 \xi_{it} + X'_{it-1}\gamma + \Psi_t + \varepsilon_{it} \quad (1)$$

where i indexes companies i=1,2..N and t indexes year t=1,2..T. n is (log) average company employment during the year, w is the (log) average wage at the company, k is the (log) capital stock. ξ is a demand shock proxy which consists of the growth in log real sales and Ψ_t represent a set of common time effects (year dummies) which will control for aggregate effects including aggregate demand.[8] ε_{it} is a serially uncorrelated but possibly heteroskedastic error term.

A number of aspects of our estimation of equation (1) warrant comment. First, financial factors, represented by the regressors X_{it-1}, are considered. Despite the large literature finding evidence of financial conditions in shaping fixed and inventory investment, there are few studies which allow for such a role in the context of labour demand models. Following from the investment literature a role for the following financial variables is also considered: cash flow CF/K, liquidity m/K, the flow borrowing ratio (ie. the ratio of interest payments to cash flow) br, and underlying net indebtedness of the firm $(B-m)/K$. In view of the analysis of Nickell and Nicolitsas (1999) we are especially interested in the role of the variable br, which they consider their preferred measure of financial pressure picking up both the premium in borrowing costs and the probability that credit is rationed.

Second, separate labour demand equations for rigid/permanent and flexible/temporary labour are estimated. This will also allow an analysis of whether and to what extent, flexible labour carries a greater burden of adjustment to demand and financial shocks than permanent employment. One issue which arises in the context of estimating the employment equation for temporary contracts is that a significant proportion of firms do not employ temporary employees. This discouraged Bentolila and Saint-Paul (1992)

[8] The demand shock variable is not considered in the analysis of Nickell and Nicolitsas (1999) but is a key consideration of Bentolila and Saint-Paul (1992) and is therefore included here. As in Bentolila and Saint-Paul (1992) we are unable to construct a relative wage of permanent/temporary labour which would ideally appear in Equation 1. We essentially assume this to be subsumed into the fixed effects.

from looking at the behaviour of (log) temporary employment explicitly. With only minor reservations, we circumvent this issue by recoding zero values on temporary employment to 0.5. This is clearly preferable to selecting out those firms that have no temporary employees. If we instead add an arbitrary constant of 0.5 to all values this provides a very similar pattern of results to that presented below.[9]

The model includes a demand shock variable, ξ_{it}, following Bentolila and Saint-Paul (1992). Support for the use of the change in log sales as a demand shock variable is provided by the fact that an AR(1) model for log real sales produces a coefficient of 0.997. We will be interested in considering how permanent and temporary labour demand varies in response to this demand shock variable.

Since during our extended sample period of 1985 to 2001, a major recession took place in Spain, the trough of which was reached in 1993, we can consider how the macroeconomic influences on company-level labour demand, in the form of the set of time effects Ψ_t vary. These can be compared when estimating separately for permanent and temporary contract employees to see to what extent temporary employment proved more responsive to the economic cycle than permanent employment during that period.

Finally, we also focus on the effects of the policy reform of 1997 which reduced payroll taxes and dismissal costs for those hired under permanent contracts described in Section 2. We add to Equation 1 a series of interaction terms between our variables

[9] Note that the issue of zero values for temporary employment does not produce a censored regression or Tobit model since there is no unobserved threshold. The problem is in taking logs. We have also estimated these models in levels, with the results being qualitatively similar (these results are available from the authors on request).

of interest and a post-1997 dummy to consider variation in response to the 1997 reform.

Our estimator consists of the GMM-System estimator proposed by Arellano and Bover (1995) and examined in detail by Blundell and Bond (1998). These models control for fixed effects with the estimator being an extension of the GMM estimator of Arellano and Bond (1991) and estimates equations in levels as well as in first-differences. Where there is persistence in the data such that the lagged levels of a variable are not highly correlated with the first difference, also estimating the levels equations with a lagged difference term as an instrument offers significant gains, countering the bias associated with weak instruments (see Blundell and Bond, 1998). We are able to further motivate the use of the GMM-System estimator by showing that the data would, under the GMM first difference estimator, suffer from this weak instruments problem. This is achieved by comparing the first difference GMM estimate of the coefficient of the lagged dependent variable to those obtained under OLS and Within Groups estimators (see Bond *et al.* 2001). Whilst an OLS estimate of the lagged dependent variable is upward biased, the Within Groups estimate is downward biased. If the estimate obtained using the first-differenced GMM estimator lies close to or below the Within Groups estimate then the likely explanation is that the first-differenced GMM estimator is subject to the weak instruments problem and the GMM-System estimator should be employed (where the latter's estimate of the coefficient of the lagged dependent variable should lie between the OLS and Within Groups estimates). We report results from these comparisons in the text below. The estimation method requires the absence of second order serial correlation in the first differenced residuals for which the test of Arellano and Bond (1991) is presented (labelled M_2). If the underlying model's residuals are indeed white noise then first-order serial correlation should be expected in the first-differenced residuals for which we also present the test of Arellano and Bond (1991), labelled M_1. We also report the results of the Sargan test for instrument validity in the GMM-system equations and the Difference-Sargan statistic, which tests the validity of the additional moment conditions associated with the levels equations.

3.1 The Data

The data employed are derived from an annual survey of non-financial firms conducted by the Central Balance Sheet Office of the Banco de España (see Banco de España, 2000). This is a large scale survey used extensively by the Bank of Spain at an aggregate level to inform its assessment of the Spanish corporate sector. In terms of gross value added the survey respondents jointly represent around 35 per cent of the total gross value added of the non-financial corporate sector in Spain. This paper employs data for the period 1985 to 2001 for which the coverage of the survey has been relatively stable. We impose restrictions that the firm has at least 10 employees, is based principally in the manufacturing sector and is present in the panel for at least five consecutive years. This produces an unbalanced sample of 3,400 manufacturing companies with between 5 and 17 annual observations per company (see Data Appendix).

Summary statistics on the main variables of interest are presented in Table 1. The survey's coverage of small firms increased during the mid-late 1980s which accounts for the decline in the average size of firms over time. The proportion of average company employment on temporary contracts increases during the sample period.

3.2 Estimation results

Our first set of estimation results is presented in Table 2. This presents results obtained for total employment. No evidence was found here of a significant role for a second lag in employment and the results we report omit this. Further justification for the use of the GMM-System estimator is obtained by comparing estimates of the coefficient on the lagged dependent variable under OLS (an upward biased 0.975), Within Groups (a downward biased 0.809) and that of the first-differenced GMM estimator (at 0.718 below the downward biased Within Groups estimate), whilst that of the GMM-System estimator lies in between the former two estimates at 0.915. This suggests that the first-differenced GMM estimator is subject to the weak instruments problem examined by Blundell and Bond (1998). In our preferred estimates reported below we selected instruments dated from $t-4$ to $t-6$ in the first-differenced equation and $\Delta t-3$ in the levels equation in order that the Sargan test statistic reported was not significant at conventional levels, although these estimates proved very similar to those where the instrument set included instruments dated $t-2$, $t-3$ and $\Delta t-1$.

The estimated long-run wage elasticity in column 2 is –0.37. This compares to an estimate of -1.86 obtained by Bentolila and Saint-Paul (1992) which they considered to be quite high relative to previous (mostly aggregate) studies. Our estimate is closer to the two other firm-level studies reviewed by Addison and Teixeira (2003), which found long-run labour cost elasticities of -0.24 for the UK (Arellano and Bond, 1991) and -0.71 for Portugal (Addison and Teixeira, 2001). Nickell and Nicolitsas (1999) obtain a wage elasticity of UK firm-level labour demand of -0.76. Our estimate for the long-run employment elasticity with respect to the demand shock is 0.05, slightly smaller than the elasticity found in Bentolila and Saint-Paul (1992) of 0.09 but identical to that obtained for the UK (Arellano and Bond, 1991) and similar to that of 0.03 for Portugal (Addison and Teixeira, 2001).

Our results indicate the presence of strong effects from financial factors to labour demand among Spanish manufacturing firms. This is one of our key results. The borrowing ratio term br_{it-1}, the measure of financial pressure favoured by Nickell and Nicolitsas (1999), designed to capture the premium on borrowing costs or the probability of credit being rationed completely, is statistically significant with a coefficient (robust standard error) of -0.055 (0.016). The addition of the cash flow term, which we include partly as a control variable, falls short of significance with a coefficient (standard error) of 0.039 (0.032), and gives rise to a minor reduction in the borrowing ratio coefficient (in absolute terms) to -0.041, but this remains significant. In column 3 we consider a liquidity variable and is not significant. Nor does the inclusion of the net indebtedness term $(B-m)/K_{it-1}$ suggest a role for debt, whilst controlling for the flow borrowing ratio term br_{it-1}. When other financial variables are omitted the $(B-m)/K_{it-1}$ term is negative but is not significant with a coefficient (robust standard error) of -0.016 (0.018). In general, the results suggest a role for financial factors in the form of the flow borrowing ratio variable consistent with the interpretation that financing constraints influence labour demand. The diagnostic tests report satisfactory values with, crucially, the absence of second-order serial correlation under the M_2 test and the Sargan tests also returning insignificant values at standard confidence levels.

In Table 3 we present estimates of labour demand equations estimated separately for permanent and temporary labour. The pattern of results confirms the hypotheses outlined previously. First, labour demand for temporary labour is more responsive to the company-specific demand shock ξ_{it}. In the first equation reported for temporary employment, this attracts a coefficient (robust standard error) of 0.628 (0.191), whilst in the similar specification for permanent employment it attracts a coefficient (standard error) of 0.112 (0.053).

Second, concerning the responses of rigid and temporary employment to an aggregate recession, in Figure 1 we plot the implied variation according the estimated aggregate time effects Ψ_t estimated in the two equations (in columns 1 and 4). It is clear that the aggregate effects estimated for permanent employment imply less aggregate flexibility than those for temporary employment. Especially notable is how the recession in 1993 was associated with a much larger aggregate negative effect for temporary labour demand.

Third, the estimates also confirm a lower degree of persistence for temporary than for permanent employment, consistent with lower adjustment costs in the case of the former. The lagged dependent variable attracts a coefficient (robust standard error) of 0.775 (0.045) in the case of temporary employment and 0.866 (0.017) in the case of permanent labour in the first specification reported for each labour type.

Fourth, we also find that the average wage elasticity of labour demand is greater for temporary than permanent employment. This finding indicates that there is scope for potential insider-outsider effects (eg. Bentolila and Dolado, 1994) since the employment consequences of wage demands are experienced more heavily by temporary employees. Indeed, in the case of permanent employment, the long-run elasticity although negatively signed is insignificantly different from zero, whilst for temporary employment it is estimated at -2.11.

Fifth, the consideration of financial factors for the two types of labour also reveals a stronger effect in the case of temporary labour, in particular in the case of the flow borrowing ratio highlighted by Nickell and Nicolitsas (1999). This term attracts an insignificant coefficient (robust standard error) of -0.027 (0.025) for permanent employment (column 1) and -0.355 (0.112) for temporary employment. The cash flow term CF/K_{it-1} is however at the margin of significance in the case of permanent labour demand but falls far short of significance in the case of temporary labour.

We conclude the discussion of these results by noting the possibility that the introduction of temporary contracts when one recognises the importance of financial constraints, may have had more positive consequences for employment than would otherwise have been the case. If temporary contracts, by reducing the 'fixity' of labour, can help employers overcome financial constraints then labour demand may increase as a result[10].

The 1997 Reform and Payroll Taxes

We now consider the impact on labour demand of the 1997 labour market reform described in Section 2 in the form of the role of payroll taxes and the variation induced by the 1997 reform.[11] The assessment of the effects of this reform is relevant since this

[10] See Rendón (2001) for a related discussion. This author shows that firms substitute temporary labour for permanent one and use less debt as their financial position improves.

[11] We also considered repeating our basic labour demand models with the addition of a set of interaction terms between the regressors of interest and a dummy variable for the post-1997 period. This approach failed to indicate any significant changes post-1997, however.

type of transitory reductions in payroll taxes has been frequently used in the Spanish case. As noted above, the impact of the 1997 reform will depend on the changes in the structure of the firm's employment. Using data on payroll tax contributions made by the firm alongside its direct wage bill we construct an average payroll tax experienced by the firm (see also Gruber, 1997). In addition to variation in tax payments induced by the policy reform, observed variation across firms will have a number of further sources. Firms will differ in the number of workers they employ with earnings above and below the earnings thresholds for contributions to be made.[12]

We measure the impact of the 1997 reform on the firm's change in the payroll tax as the 1998 average payroll tax less its value in 1996.[13] The mean payroll tax over the full period is 27.8 per cent calculated in this way, which compares to a figure in 2000 of 30.6 per cent available from aggregate figures for Spain. (Source: the Spanish national earnings survey, '*Indice de Costes Laborales*'). The mean change associated with the 1997 reform is a reduction in the payroll tax of 0.5 percentage points but, importantly, with significant variation across firms. The 25th percentile firm experiences a change of -1.2 percentage points whilst the 75th percentile is +0.7 percentage points.

[12] These lower and upper thresholds state that if the monthly income of an employee lies below (above) a given amount, social security contribution are liable as if income was at that threshold. In 2002, for instance, the lower (upper) threshold was 516 (2,574.9) Euros. Previously, the thresholds have been occupation-specific.

[13] Thus $\Delta t = \frac{t_2 w_2 N_2}{w_2 N_2} - \frac{t_1 w_1 N_1}{w_1 N_1}$, where twN represents total social security contributions by the firm and wN are the direct remuneration costs. Δt is the change in average payroll tax experienced by the firm calculated between 1996 and 1998. The variable is not time-varying for each firm. Since information on social security contributions is not available for all companies, its use involves a reduction in the sample size by 501 firms.

In a competitive labour market and with an inelastic labour supply curve any change in payroll taxes would induce an offsetting response to gross pay leaving wage costs and employment unchanged. However, in non-competitive labour markets where there is some element of real wage resistance, some of the burden of these costs can be shifted onto the employer resulting in employment effects. The extent to which these effects are important remains unsettled. Most macroeconomic studies reviewed in Nickell and Layard (1999) support their existence but that their magnitude is modest. However, Daveri and Tabellini (2000) provide evidence suggesting these effects are large, at least in continental Europe.[14] There has been very little microeconomic evidence on this subject however. Gruber (1997) provides company-level evidence based on a panel of firms from Chile where the sample period includes a very pronounced change in payroll tax rates. The evidence there suggested complete wage shifting to the payroll tax change with no effect on employment.[15]

Table 4 presents analysis of the effects of the policy reform based on our payroll tax variable, Δt. Since the policy reform affected permanent employment directly, we concentrate on the permanent and total labour demand equations. Since firms responded to social security subsidies by changing the composition of their labour force in favour of permanent jobs, thereby reducing their own tax burden, it is clear that our measure of the firm's change in the payroll tax is not exogenous. To address this

[14] Pissarides (1998) studies how these effects differ under a number of different labour market models and how the effects interact with the structure (progressivity) of the tax system and unemployment benefits.

[15] See also Kugler and Kugler (2002) for evidence on Colombia.

endogeneity issue we use the change in the average sectoral tax rate as an instrument.[16] Our payroll tax variable enters the labour demand equation negatively signed and statistically significant (albeit at the margin of significance in the case of permanent rather than total employment) indicating that the *reduction* in payroll taxes associated with the 1997 policy reform increased employment. The fact that the result is only significant in the case of permanent workers is perhaps not too surprising since the reform only affected new permanent contracts. The point estimate implies that a 5 percentage point reduction in the payroll tax increases labour demand by 8 per cent. This estimate is of the same order as that in Pissarides's (1998) simulations. In the case of holding unemployment benefits fixed in real terms, Pissarides's (1998) estimates for his union bargaining model imply that a 5 percentage point reduction in payroll taxes would reduce equilibrium unemployment by around 1 percentage point. Nickell and Layard (1999) suggest effects a little larger than this (ie. around 13 per cent or 1.3 percentage points) as typical of the time-series evidence but are subject to "a great deal of uncertainty" while the estimates of Daveri and Tabellini (2000) for the continental European economies are somewhat larger still.

4 Conclusion

This paper has examined the demand for labour in a highly regulated labour market and contrasted the demand for flexible and rigid labour. It has previously been argued that the case of Spain is especially informative regarding labour demand and certain policy interventions designed to increase flexibility in a market where many workers

[16] See Data Appendix for the precise definition of this instrument.

enjoy high levels of employee security. Relevant characteristics of the Spanish labour market include the significant differences in protection rights between permanent and temporary contract labour, the high incidence of temporary contracts over a significant period of time, and the significant policy reforms that have been introduced affecting such contracts. In this paper we have extended this literature in a number of ways.

First, we have explicitly compared estimates of the demand for flexible labour with that for labour employed under permanent contracts. Second, we have considered a role for financial factors. Despite a large number of studies of the role of financial conditions in affecting other factor demands, notably fixed investment, previous studies of labour demand have largely ignored the possible influence of financial factors. Third, we have examined a significant policy reform in 1997. Such reforms provide the ideal context to appraise arguments concerning the effects of labour market regulation. Moreover, the 1997 policy was different in nature to the earlier 1984 reform responsible for the initial liberalisation and subsequent widespread adoption of temporary contracts. Rather than offer an alternative to permanent contracts, the 1997 reform focused on the reduction in payroll taxes associated with permanent contracts themselves. The appraisal of this policy reform, which as emphasised by Kugler *et al.* (2002) was not a marginal one, is of further significance since in 2001 its scope was further extended. Fourth, we have revisited the empirical predictions put forward by Bentolila and Saint-Paul (1992) who focused on variations in labour demand over the cycle as a result of temporary contracts. In this context our ability to extend the time dimension of a panel of Spanish firms, not least to include a period of recession, was highlighted.

Our results make a number of contributions. We have estimated how and to what extent the demand for flexible labour responds differently from that for rigid labour. As theory suggests, the demand for flexible labour is less persistent, more responsive to wage costs and is more sensitive to financial factors. This leads to the conclusion that where adjustment in terms of employee numbers is required, the burden of such adjustment is borne disproportionately by those with temporary contracts. The notion that such contracts thereby enhance the ability of employers to undertake adjustments to a number of factors is clearly borne out by the data.

Relative to the predictions confronted with a panel of Spanish manufacturing firms by Bentolila and Saint-Paul (1992), we provide supportive evidence of a number of predictions that Bentolila and Saint-Paul (1992) were unable to uncover since their study was restricted to a limited time dimension (from 1985 to 1988). This includes the finding that flexible labour demand is more sensitive to the demand shock and to the aggregate economic cycle than that of permanent labour.

As part of our assessment of the 1997 reform we also considered the adjustment of payroll taxes, exploiting company-specific information in payroll costs. Our estimates suggest that a 5 percentage point reduction in payroll taxes increases employment by 8 per cent. Two notes of caution are in order: First, it should be noted that the introduction of social security subsidies (as in the 1997 reform) have been usually implemented on a transitory basis and it is uncertain whether the effects on employment dissapear once the policy measure is reversed Second, our analysis ignores the revenue losses associated to the reform that should be taken into account for an overall assessment of the reform. Our evidence adds to the largely macroeconomic-based evidence of the effects of payroll taxes. Since this aggregate evidence has produced such a wide range of estimates, microeconomic evidence of this kind may prove particularly useful in informing the debate concerning the effects of payroll taxes on employment.

Table 1: Summary Statistics

		1985-88	1989-92	1993-96	1997-2001	1985-2000
N	employment	324.562	265.042	199.795	187.173	240.794
$(N^T/N)*100$		9.187	14.984	19.061	19.686	16.018
Δw	wage growth	0.023	0.041	0.013	0.012	0.023
br	borrowing ratio	0.347	0.414	0.402	0.241	0.352
CF/K	cash flow	0.168	0.130	0.120	0.158	0.142
$(B-m)/K$	net indebtedness	0.280	0.240	0.241	0.216	0.243
m/K	liquidity	0.110	0.095	0.086	0.100	0.097
ξ	real sales growth	0.039	-0.018	0.027	0.044	0.022
Y	real sales (1995 prices)	45,943.4	39,278.95	36,636.84	44,639.01	41,389.46
observations		6,858	8,918	8,336	8,373	32,485

Notes: Tables reports sample means. Real sales are thousands of euros (1995 prices).

Table 2: Total labour demand

	[1]	[2]	[3]	[4]	[5]
n_{it-1}	0.915 (0.013)	0.919 (0.012)	0.919 (0.012)	0.923 (0.011)	0.925 (0.011)
k_{it}	0.042 (0.008)	0.043 (0.007)	0.041 (0.008)	0.039 (0.007)	0.037 (0.007)
Δw_{it}	-0.520 (0.069)	-0.503 (0.061)	-0.516 (0.069)	-0.475 (0.059)	-0.512 (0.063)
w_{it-1}	-0.029 (0.021)	-0.031 (0.020)	-0.027 (0.021)	-0.018 (0.019)	-0.015 (0.020)
br_{it-1}	-0.055 (0.016)	-0.041 (0.019)	-0.057 (0.016)	-0.047 (0.019)	-0.067 (0.017)
CF/K_{it-1}		0.039 (0.032)		0.048 (0.030)	
m/K_{it-1}			-0.005 (0.032)		0.016 (0.039)
$(B-m)/K_{it-1}$				0.019 (0.018)	0.024 (0.024)
ξ_{it}	0.179 (0.038)	0.180 (0.035)	0.193 (0.036)	0.173 (0.032)	0.186 (0.034)
year effects	yes	yes	yes	yes	yes
M_1 (p-value)	0.000	0.000	0.000	0.000	0.000
M_2 (p-value)	0.143	0.147	0.184	0.132	0.142
Instruments	t-4..t-6, Δt-3	t-4..t-6, Δt-3	t-4..t-6, Δt-3	t-4..t-6, Δt-3	t-4..t-6, Δt-3
Sargan (p-value)	0.237	0.169	0.299	0.196	0.324
Difference-Sargan (p-value)	0.184	0.373	0.163	0.353	0.175
companies	3,400	3,400	3,400	3,400	3,400
observations	29,085	29,085	29,085	29,085	29,085

Notes: Estimation by GMM-SYSTEM estimator using the robust one-step method (Blundell and Bond, 1998; Arellano and Bond, 1998). Sargan is a Sargan Test of over-identifying restrictions (p-value reported), distributed as chi-square under the null of instrument validity. Difference-Sargan is a Sargan Test of the validity of the additional moment conditions associated with the levels equations (p-value reported), distributed as chi-squared under the null of instruments validity, M_j is a test of jth-order serial correlation in the first-differenced residuals. These are both distributed as standard normals under the null hypotheses. Asymptotic robust standard errors reported in parentheses. Instruments used are the lagged values of the regressors dated from t-4 to t-6 in the first-differenced equation and the first difference of the regressors dated t-3 in the levels equation.

Table 3: Permanent and Temporary labour demand

	Permanent employment			Temporary employment		
	[1]	[2]	[3]	[4]	[5]	[6]
n_{it-1}	0.866 (0.017)	0.868 (0.017)	0.865 (0.019)	0.775 (0.045)	0.774 (0.039)	0.783 (0.041)
k_{it}	0.077 (0.012)	0.081 (0.015)	0.080 (0.018)	0.151 (0.042)	0.155 (0.040)	0.142 (0.039)
Δw_{it}	-0.345 (0.102)	-0.296 (0.093)	-0.353 (0.113)	-1.097 (0.407)	-1.042 (0.366)	-0.965 (0.392)
w_{it-1}	-0.002 (0.039)	-0.007 (0.042)	-0.004 (0.042)	-0.476 (0.172)	-0.471 (0.162)	-0.405 (0.162)
br_{it-1}	-0.027 (0.025)	-0.002 (0.028)	-0.044 (0.024)	-0.355 (0.112)	-0.341 (0.130)	-0.310 (0.112)
CF/K_{it-1}		0.085 (0.054)			0.001 (0.223)	
m/K_{it-1}			-0.063 (0.079)			0.187 (0.196)
ξ_{it}	0.112 (0.053)	0.121 (0.046)	0.165 (0.050)	0.628 (0.191)	0.696 (0.182)	0.589 (0.166)
year effects	yes	yes	yes	yes	yes	yes
M_1 (p-value)	0.000	0.000	0.000	0.000	0.000	0.000
M_2 (p-value)	0.420	0.435	0.510	0.906	0.877	0.919
Instruments	t-4..t-6, Δt-3	t-4..t-6, Δt-3	t-4..t-6, Δt-3	t-4..t-6, Δt-3	t-4..t-6, Δt-3	t-4..t-6, Δt-3
Sargan (p-value)	0.085	0.058	0.140	0.202	0.146	0.314
Difference-Sargan (p-value)	0.098	0.169	0.241	0.126	0.119	0.306
companies	3,400	3,400	3,400	3,400	3,400	3,400
observations	29,085	29,085	29,085	29,085	29,085	29,085

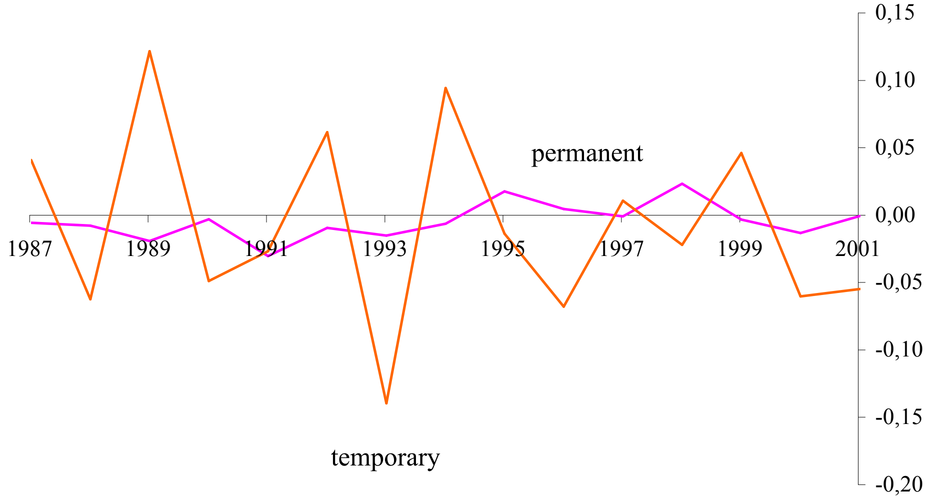
Notes: See Table 2.

Table 4: The 1997 Reform and Payroll Taxes

	Total employment	Permanent employment
n_{it-1}	0.944 (0.011)	0.917 (0.016)
k_{it}	0.034 (0.008)	0.063 (0.013)
Δw_{it}	-0.567 (0.075)	-0.469 (0.110)
w_{it-1}	-0.041 (0.024)	-0.049 (0.042)
br_{it-1}	-0.067 (0.017)	-0.060 (0.024)
ξ_{it}	0.193 (0.046)	0.148 (0.057)
Δt_{it}	-0.006 (0.004)	-0.016 (0.008)
year effects	yes	yes
M_1 (p-value)	0.000	0.000
M_2 (p-value)	0.129	0.561
Instruments	t-4..t-6, Δt-3	t-4..t-6, Δt-3
Sargan (p-value)	0.096	0.458
Difference-Sargan (p-value)	0.024	0.052
companies	2,899	2,899
observations	23,534	23,534

Notes: See Table 2. We add the change in the average tax rate (at the sectoral level) to the list of instruments.

Figure 1: Aggregate Effects on Permanent and Temporary Employment



Data Appendix

Table A.1 tabulates the number of time-series observations per company.

Table A.1: Panel structure

No of records	5	6	7	8	9	10	11
Companies	555	471	411	301	245	167	161
No of records	12	13	14	15	16	17	Total
Companies	182	150	178	202	112	245	3,400

Employment

The number of employees during year. The data also distinguish between the number on permanent and temporary contracts.

Wages

The average company wage is given by direct employment costs (not including social security contributions) dividend by employment head count and deflated by the GDP deflator.

Payroll taxes

The payroll tax paid by the firm is equal to $t = \frac{SS}{wN}x100$, where SS represents total social security contributions by the firm and wN are the direct remuneration costs.

The change in the payroll tax associated with the 1997 reform is calculated as the difference in the company payroll tax between 1996 and 1998. This value is then kept by that firm for the period 1997 to 2001 or whenever the firm leaves the panel and is equal to zero before 1997.

This variable is instrumented with the difference in the sectoral average payroll tax rates between 1996 and 1998, using a sectoral disaggregation in 39 industries. (Source: Spanish National Accounts. Base 1995. Production and generation of income accounts by industry). As in the case of the firm-level measure, this value is then kept by that sector for the period 1997 to 2001 and is equal to zero before 1997.

Liquidity (m/K)

Liquid assets are given by cash and equivalent, normalised on capital stock.

Debt (B/K)

Total outstanding debt divided by capital stock, K (see below). Net debt $(B-m)$ subtracts cash and equivalent from the numerator.

Capital stock (K)

This is given by the sum of fixed assets at replacement cost (calculated by the Central de Balances (CBA) of the Bank of Spain) and working capital less provisions.

Cash flow (CF)

Post-tax profit plus depreciation of fixed assets.

Borrowing ratio (br)

Interest payments divided by cash flow. Where companies have a negative value for the denominator their borrowing ratio is set equal to 1.

Real Sales (S)

Total company sales, deflated by the GDP deflator.

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