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THE VALUE OF FLEXIBLE CONTRACTS: EVIDENCE FROM AN ITALIAN PANEL OF INDUSTRIAL FIRMS

By Piero Cipollone* and Anita Guelfi**

Abstract

Since the mid-1980s fixed-term contracts have been used in many European countries to reduce firing costs. As this strategy may have led to segmented labour markets, recent policy interventions have enhanced permanent jobs by cutting their labour costs. Efficient design of these policies requires knowledge of the costs associated with employment protection legislation. In this paper we evaluate these costs by measuring firms' willingness to trade fixed-term for open-ended contracts in exchange for a cut in the labour cost of permanent jobs. Our results are based on a panel of Italian firms in the engineering sector whose labour costs were reduced by a tax credit granted to firms hiring workers on open-ended rather than fixed-term contracts. The trade-off is identified by comparing how the composition of recruitment by type of contract changed for firms that received the tax credit and those that did not. Potential distortions due to self-selection into the programme, firm-specific time-varying shocks or mechanical correlation induced by the selection rule into the programme, are accounted for by estimating the spurious effect of the tax credit in the years when it was not in force. Estimation is carried out in both a parametric and non-parametric setting that uses p-score to control for different probabilities of receiving the tax credit. We found that firms value the possibility of hiring one per cent new workers on a fixed-term contract as much as a cut in the labour cost of an open-ended worker in the range of 1.3-2.8 per cent. This result helps to explain recent employment growth in Italy, where the share of fixed-term contracts among new hires grew from 34 to 42 per cent between 1995 and 2003. Using our most conservative results, we evaluate that the labour cost reduction associated with this expansion amounted to anything between 10.4 and 22.4 per cent. Given the elasticity of employment to wages, the advent of flexibility in the Italian labour market can account for a large share, between 37 and 80 per cent, of employment growth in the private sector.

JEL classification: D78, H25, J23, J38

Keywords: tax credit, open-ended contracts, fixed-term contracts, firing costs.

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1. Introduction[1]

How large are the costs born by firms because of employment protection legislation (EPL)? This question is at the core of the renewed policy effort of the European Union, as confirmed recently in the Kok report (EU, 2003), towards "flex-security", that is the attempt to combine better and more secure jobs with a highly flexible labour market. The basic idea is to increase the number of permanent jobs by cutting their labour costs, thereby compensating firms for giving up the flexibility associated with fixed-term contracts. The efficient design of this policy requires knowledge of the costs associated with EPL.

Recent examples of such a policy can be found in both Spain and Italy. In 1997, Spain drastically cut payroll taxes on new permanent contracts for a period of two years in an attempt to reduce the segmentation of the Spanish labour market induced by the 1984 liberalization of temporary contracts. Firing costs for unfair dismissal were also lowered by around 25 per cent (Benito and Hernando, 2003). Kugler, Jimeno and Hernanz (2002) estimated that this reform reduced the cost of hiring young workers on a permanent contract by about 10 and 7 per cent in first and second year respectively. In the year 2000, Italy adopted a similar provision by granting a large tax credit to firms hiring workers on an open-ended contract. The implied cut in labour costs has been evaluated at between 9 per cent and nearly 60 per cent depending on both the industry and the geographical area (Cipollone and Guelfi, 2003).

Knowledge of the value firms attach to flexible contracts is crucial for an optimal design of this type of compensation policy. Despite the simplicity of the underlying concept, this information is rather difficult to obtain because of the multidimensional nature of the costs associated with firing. Along with monetary expenditures (severance payment), there are burdens associated with the length of the administrative and legal procedures and the cost of uncertainty. OECD (1999) provides a complete list of the costs generated by employment protection legislation.

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Three main strands of literature have tried to measure the magnitude of firing costs. A first line of research has attempted to evaluate these costs within the framework of dynamic labour demand,[2] which was developed after the seminal work of Oi (1962). Along this line Rota (2004) estimated that in Italy these costs account for as much as 15 per cent of monthly wages. In a similar setting Aguirregabiria and Alonso-Borrego (1999) found that firing costs in Spain amounted to 51 per cent of the gross annual wages of a permanent worker over the period 1982-1993. For France Goux, Maurin and Pauchet (2001) found that the bulk of the adjustment cost is due to the firing of permanent workers.

A second strand of literature has attempted to compute directly the cost of EPL by examining the regulation or the actual costs declared by firms. Using direct evidence, Abowd and Kramarz (2003) evaluate that in 1992 separation costs in France amounted to anything between 56 and 126 per cent of average labour costs. Garibaldi and Violante (2005) suggest that in Italy firing costs are equivalent to about 18 monthly wages.

A third group of empirical studies has attempted to quantify the size of firing costs by estimating their impact on the level of employment. This large body of literature has not reached any conclusive consensus (OECD, 1999, 2006).

In this paper we offer and alternative way to ascertain these costs by evaluating the willingness of firms to give up flexibility in exchange for a cut in labour costs. We estimate the labour demand for fixed-term contracts compared with that for open-ended contracts as a function of the relative wage and firm specific controls and derive the money value that firms attach to fixed-term contracts. Hence, our framework allow us to infer the value of flexibility directly rather than from ex-post computation. However, the reliability of this approach impinges on the assumption that the relative wage coefficient in the estimated labour demand equation measures a causal effect. In most settings this is a heroic assumption because of firms' unobservable heterogeneity and because the standard response of labour supply to changes in relative wages acts as a confounding factor. In this paper we overcome these problems by resorting to a policy-induced shift in labour demand not directly linked to a variation in wages. To this end we exploit the introduction in Italy, at the end of 2000, of a tax credit for firms choosing to hire workers on open-ended rather than fixed-term contracts. This regulation created a trade-off for the firm between fewer flexible contracts and a substantial wage cut. Observing the variation in firms' response, we are able to uncover the rate at which this trade-off occurs. We run our exercise on a panel of about 310 Italian firms

[2] Hamermesh and Pfann (1996) have recently reviewed the literature on the nature and determinants of adjustment costs.

operating in the engineering sector in the period 1998-2002, which hired both fixed-term and open-ended workers for the whole period.

Identification of the effect of the tax credit on the relative demand for fixed-term and open-ended workers is based on the comparison of changes in this relative demand of firms that received the tax credit and those that did not. Recovering the causal effect means controlling for potential distortions due to self-selection into the programme because of firm time-specific shocks. On top of this standard problem, we need to take into account that the selection rule into the tax credit – raising the stock of permanent workers – might be correlated with the relative labour demand. However, neither of these problems affects the estimate of the causal effect of tax credit on relative demand as long as potential bias is constant over time. In the standard diff-in-diff setting this time invariance is assumed away. In our setting we can instead estimate the bias because we observed both relative labour demand and selection rule in the years in which the tax credit was not in force. As this bias is always zero in our data, we recover the causal effect a standard difference in means estimator, using both a parametric setting and the non-parametric version devised by Heckman et al. (1997), Heckman et al. (1998) and Blundel and Costa Diaz (2000).

We estimate that hiring one per cent new workers on a fixed-term contract is worth as much as a cut of between 1.3 per cent (parametric setting) and 2.8 per cent (non-parametric setting) in the labour cost of a worker hired on an open-ended contract. The estimate suggests that the effect of the tax credit was large. It implies that in the period 1995-2003 firms enjoyed an overall labour cost reduction ranging from 10.4 to 22.4 per cent as the share of fixed-term contracts out of all newly hired workers rose from 34 to 42 per cent. In Italy the long-run elasticity of employment to wages is estimated to be around -0.3; therefore we calculated that the rise in the share of flexible contracts can explain a large part, from 3.1 to 6.7, of the 8.4 percentage point increase in private sector total employment.

This paper is organised as follows. In Section 2 we discuss our empirical strategy while Section 3 is devoted to a detailed illustration of the characteristics of the tax credit; Section 4 presents the data; Section 5 discusses the identification of the causal effect of the tax credit on relative labour demand; Section 6 presents the main results; Section 7 recovers the trade-off between flexibility and wage cut. Section 8 uses this result to evaluate the effect of the increase in fixed-term contracts among newly hired workers in Italy between 1995 and 2003. Section 9 concludes.

2. The statistical framework for estimating the value of fixed-term contracts

In this paper we define the value of flexibility as the semi-elasticity :

$$VFC = \frac{\frac{\Delta w}{w}}{\Delta\left(\frac{H^{oe}}{H^{ft} + H^{oe}}\right)}$$

where w represents the labour cost of an open-ended contract, H indicates recruitments and superscripts ft and oe indicate fixed-term and open-ended contracts, respectively. In words, our definition of the value of flexibility, measured in terms of the labour cost of an open-ended worker, represents the percentage reduction in the labour cost w that firms are willing to trade for an increase of one percentage point in the share of total new workers hired on an open-ended contract.

In order to measure how much firms value flexible contracts, we use a simple model that relates the ratio between the shares of fixed-term compared with open-ended engagements to their corresponding relative wages. The workhorse used in this paper is a standard partial equilibrium model for labour demand and the supply of fixed-term relative to open-ended workers:[3]

(1) $\log(\frac{H^{ft}}{H^{oe}})^{D}_{it} = \beta_0 + \beta_1 \log(\frac{W^{ft}}{W^{oe}})_{it} + \beta_2 TC_{it} + \beta_3 Z_{it} + \nu_i + \varepsilon_t + \eta_{it}$

(2) $\log(\frac{H^{ft}}{H^{oe}})^{S}_{it} = \gamma_0 + \gamma_1 \log(\frac{W^{ft}}{W^{oe}})_{it} + \gamma_2 SH_{it} + \gamma_3 Z_{it} + \phi_i + \lambda_t + \mu_{it}$

where $\log(\frac{H^{ft}}{H^{oe}})^{S}_{it}$ and $\log(\frac{H^{ft}}{H^{oe}})^{D}_{it}$, which represent the (log of) supply and demand for fixed-term relative to open-ended workers (for firm i at time t), depend on the (log of) relative labour cost of the two types of workers ($\log(\frac{W^{ft}}{W^{oe}})_{it}$), on Z_{it}, that is the firm's time-varying characteristic, on TC_{it} and SH_{it}, two shifters specific for demand and supply respectively, on ν_i and ϕ_i standing for time invariant firm's characteristics, on year effects (ε_t and λ_t), and on idiosyncratic shocks (η_{it} and μ_{it}). The labour demand shifter TC_{it} is the

[3] This simple model has been widely applied in the literature on Skill Biased Technical Changes to explain the relationship between the relative supply and relative wage of skilled and unskilled workers (Autor, Katz, Kearney 2005).

reduced labour cost induced by the tax credit $TC_{it} \equiv \log(\frac{W_{it}^{oe} - VTC_{it}}{W_{it}^{oe}})$, where VTC_{it} is the euro value of the tax credit (more on this in the next section) and W_{it}^{oe} is the labour cost of an open-ended worker. In this setting the share of open-ended contracts on total hiring is $\left(\frac{H^{oe}}{H^{ft} + H^{oe}}\right) = \frac{1}{k_{it}(1 - \frac{vtc_{it}}{w_{it}^{oe}})^{\beta_2} + 1}$ where k_{it} is the level of $\frac{H^{ft}}{H^{oe}}$ in the absence of the tax credit;[4] change in the quantity due to the labour demand shifter is $\Delta\left(\frac{H^{oe}}{H^{ft} + H^{oe}}\right) = \frac{1}{k_{it}(1 - \frac{vtc_{it}}{w_{it}^{oe}})^{\beta_2} + 1} - \frac{1}{k_{it} + 1}$ and the value of flexible contracts is

$$VFC = \frac{\frac{vtc_{it}}{w_{it}^{oe}}}{\Delta\left(\frac{H^{oe}}{H^{ft} + H^{oe}}\right)}$$

Solving for equilibrium conditions we have the following two reduced forms:

$$(3)\quad \log(\frac{H^{ft}}{H^{oe}})_{it} = b_0 + b_1 TC_{it} + b_2 SH_{it} + b_3 Z_{it} + \tilde{v}_i + \tilde{\varepsilon}_t + \tilde{\eta}_{it}$$

$$(4)\quad \log(\frac{W^{ft}}{W^{oe}})_{it} = c_0 + c_1 TC_{it} + c_2 SH_{it} + c_3 Z_{it} + \tilde{\phi}_i + \tilde{\lambda}_t + \tilde{\mu}_{it}$$

Where $b_1 = \frac{\beta_2\gamma_1}{(\gamma_1 - \beta_1)}, c_1 = \frac{\beta_2}{(\gamma_1 - \beta_1)}, b_2 = -\frac{\beta_1\gamma_2}{(\gamma_1 - \beta_1)}, c_2 = -\frac{\gamma_2}{(\gamma_1 - \beta_1)}, a_3 = \frac{\beta_3\gamma_1 - \beta_1\gamma_3}{(\gamma_1 - \beta_1)}, c_3 = \frac{\beta_3 - \gamma_3}{(\gamma_1 - \beta_1)}$.

If we had both the relative demand and relative supply shifters we would be able to identify both labour demand and labour supply elasticities, thereby pinning down firms' trade-off between flexibility and labour cost reduction. The elasticity of relative labour demand would

[4] This follows from the fact that according to equation (1) $\frac{H^{ft}}{H^{oe}} = K_{it}(1 - \frac{vct_{it}}{w_{it}^{oe}})^{\beta_2}$ and from the fact that $\left(\frac{H^{oe}}{H^{ft} + H^{oe}}\right) = \frac{1}{\frac{H^{ft}}{H^{oe}} + 1}$

be identified as $\beta_1 = \frac{b_2}{c_2}$, the elasticity of labour supply would be recovered as $\gamma_1 = \frac{b_1}{c_1}$, and the crucial parameter β_2 as $\beta_2 = \frac{b_1(\gamma_1 - \beta_1)}{\gamma_1}$.[5]

At this stage of our research we are not able to provide a credible labour supply shifter, so that we cannot identify β_1. However, a labour demand shifter may still allow us to recover the labour supply slope γ_1, thereby making some statements about β_2, the measure of the reaction of relative labour demand to the variation in its shifter. Note that when γ_1 is very large, i.e. the elasticity of labour supply to relative wages is large, then b_1 is a good approximation of β_2. In contrast, if labour supply is rigid, b_1 is lower than β_2 as it incorporates the wage reaction needed to increase the relative labour supply in order to match the higher relative labour demand. Thus b_1 represents the market reaction to the shift in labour demand and can be interpreted as the lower bound of the reduction in the share of fixed-term contracts firms are willing to trade for the labour cost cut granted by the tax credit. Note that we can obtain this important parameter by estimating equation 3 by itself. If we use b_1 rather than β_2 to compute the VFC, then this value represents the wage variation at which the market, not just the firms, trades one percentage point of fixed-term hires. This value is larger than that computed using β_2, since part of the wage rise is needed to increase the labour supply of open-ended workers. In this sense it represents an upper bound to the value for firms of flexible contracts.

The labour demand shifter we use in this paper is a recent policy intervention in the Italian labour market that reduced the relative labour cost of open-ended workers. Starting from October 2000 the Italian government rewarded firms choosing to hire workers under open-ended contracts through a tax credit of about € 413 (€ 620 for workers in the South) per month and per worker from the hiring moment until the end of December 2003[6]. This tax credit could be claimed against any type of taxes, such as income tax, social security contributions, and value-added tax.

[5] The two shifters play the role of exclusion restrictions in the structural form (1) and (2). In the absence of such restrictions we are unable to tell whether the sample variation is attributable to changes in the demand or supply schedule, as both schedules react to the same determinants. In contrast, if we have a variable such as SH that shifts only labour supply and leaves labour demand unaffected we can identify of the slope of this second schedule. By the same token we can identify the labour supply slope if we have a labour demand shifter.

[6] A more thorough description of the regulatory aspects of this recent provision is discussed in the following section.

This shifter can provide more information on the value of flexible contracts under a more restrictive assumption. As a matter of fact, if we could claim that $\beta_2 = -\beta_1$ then the slope of relative demand could be recovered as $\beta_2 = -\beta_1 = \dfrac{\gamma_1 b_1}{\gamma_1 - b_1}$; this restriction seems to be reasonable in our setting as it implies that firms' relative labour demand reacts in the same way to a change in the relative wage as to a change in non-wage labour costs. This assumption is also embedded in any model in which firms equate labour productivity with total labour cost rather than just the wage. As an example assume, that the production function is a CES

$$(5)\ Y_{it} = \left[\omega_{it}(a_t FT_{it})^{\rho} + (1-\omega_{it})(b_t OE_{it})^{\rho}\right]^{\frac{1}{\rho}}$$

where FT_{it} and OE_{it} represent new hiring under fixed-term and open-ended contracts respectively, then the relative demand for fixed-term relative to open-ended workers turns out to be, in logarithmic form, equal to

$$(6)\ \log\left(\frac{FT_{it}}{OE_{it}}\right) = -\sigma D_{it} - \sigma \log\left(\frac{LC_{it}^{ft}}{LC_{it}^{oe}}\right)$$

with $D_{it} = -\log(\omega_{it}/1-\omega_{it}) - \rho\log(a_t/b_t)$, $\sigma = 1/1-\rho$, while LC_{it} stands for the total labour cost borne by firm *i* at time *t*, *ft* for fixed-term and *oe* for open-ended. In this simple framework, σ is the elasticity of substitution and measures the number of fixed-term contracts firms are willing to trade for permanent jobs in exchange for a reduction in their relative costs. If we decompose total labour costs into their wage and non-wage components and introduce the tax credit as a non-wage cut, expression (6) becomes

$$(7)\ \log\left(\frac{FT_{it}}{OE_{it}}\right) = -\sigma D_{it} - \sigma\log\left(\frac{E_{it}^{ft}}{E_{it}^{oe}}\right) - \sigma\log\left(\frac{1+NW_{it}^{ft}}{1+NW_{it}^{oe}}\right) - \sigma\log\left(\frac{E_{it}^{oe}\left(1+NW_{it}^{oe}\right)}{E_{it}^{oe}\left(1+NW_{it}^{oe}\right)-VTC_{it}}\right)$$

with E_{it} and NW_{it} representing wage and non-wage costs for firm *i* at time *t*; VTC_{it} is the actual reduction in labour costs due to the tax credit. This equation can be mapped into equation (1) by assuming that $-\beta_1 = \beta_2 = \sigma$, $TC_{it} = -\log\left(\dfrac{E_{it}^{oe}\left(1+NW_{it}^{oe}\right)}{E_{it}^{oe}\left(1+NW_{it}^{oe}\right)-VTC_{it}}\right)$, that $\log\left(\dfrac{1+NW_{it}^{ft}}{1+NW_{it}^{oe}}\right)$ and D_{it} can be modelled with a combination of firms' time-varying controls plus fixed effects for firm and year.

3. Subsidy for open-ended contracts

Like many other OECD countries, Italy has attempted to reduce the negative effects of fixed-term contracts.[7] The strategy adopted sought to increase the mobility out of fixed-term contracts by providing fiscal incentives to firms that either transform temporary into permanent positions or directly hire workers under open-ended contracts. There are several examples of this strategy.[8] However, until 2000 these incentives were small and often targeted to particular areas, firm types or worker categories.

The Italian Finance Law for 2001 (issued at the end of 2000) provided instead a new incentive in the form of a general, automatic and quite generous tax credit to all firms hiring workers on open-ended contracts. In particular, this provision stated that every firm hiring a new worker on a permanent basis would be rewarded with a tax credit of about € 413 (€ 620 for workers in the South) per month and per worker from the moment of hiring until the end of December 2003.[9] This new tax credit applied to all hires taking place from October 2000. Thus, for a southern worker hired in October 2000 and retained until December 2003 each firm could receive about € 24,200. The tax credit was awarded only if both workers and firms met the required conditions. Workers were required to be at least 25 years old and not working with an open-ended contract in the 24 months before the hiring. Firms could apply for the tax credit if the newly hired worker raised the overall level of permanent employment – at the firm level – above the average recorded in the period between October 1999 and September 2000. The tax credit could be claimed against any kind of taxes, such as income tax, social security contributions, or value-added tax. Furthermore, it could be passed on to different fiscal years. Last, but not least, the tax credit could be cumulated with other existing subsidies.

These rules were in force until an important regulatory change was introduced in the summer of 2002. Indeed, in July 2002 the Italian Government introduced a ceiling of about € 652 million for the resources available for the new employment bonus. Since this ceiling had already been reached at the beginning of July, the tax credit was suspended. At the end of September 2002, the Government intervened again on this issue. It was decided that firms

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[7] Cipollone and Guelfi (2003) provide a full review of figures and regulations on fixed-term contracts in Italy.

[8] Examples of such attempts are the incentives to transform training-employment contracts into permanent ones or the tax credit for small firms hiring permanent workers in economically depressed areas.

[9] The contribution provided by this subsidy was large: the percentage reduction in per-capita labour costs induced by the tax credit (using data for 2000) ranges from about 9.3 per cent in the banking sector in the central and northern regions of the country to almost 60 per cent in the agricultural sector in the South. On average, in the private non-farm sector the reduction amounts to about 30 per cent in the South and 16 per cent in the central and northern regions.

would receive a tax credit up to a given ceiling of employment growth and that all credits due for the period July-December 2002 should be claimed in 2003 and by instalments. The subsidy for hires taking place in 2003 was to be regulated by the Financial Law, which simply extended the new September rules to 2003 for all firms already benefiting from the tax credit. Moreover, it prolonged the functioning of the employment bonus up to 2006, although reducing significantly the monthly amounts granted.

The new tax credit seems to have been very successful in both 2001 and 2002. According to the figures collected by the Ministry of Finance it involved on average about 110,000 workers in 2001 and 300,000 in 2002. Because of the new regulation the programme was less popular from 2003 onwards.

4. Data description

Estimation has been carried out on the micro-data collected by Federmeccanica (the Italian federation of private engineering firms) for its annual national survey on the situation of the engineering industry. Designed in 1976 to fulfil the information duties agreed with the trade unions in the collective agreement signed in that year, this survey provides yearly data on a wide set of variables covering different aspects at the firm and plant level.[10] Though the survey has experienced several changes over time in the topics chosen and/or in the detail required, information on the structure and dynamics of both employment and average earnings is available (although with different in-depth possibilities) from the very first year of data collection. Nonetheless, micro-data in electronic form are only available starting from the second half of the 1990s. We have access to data for the years 1998, 1999, 2001 and 2002 (no survey was carried out in 2000). The number of answering firms ranges from a minimum of 2448 in 2001 to a maximum of 2979 in 1998. However, we restricted the sample to firms with only a single plant that have answered at least both the 2001 and 2002 surveys, with non-missing observations for the share of hirings and for wages.[11] Moreover,

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[10] The survey provides information on the structural characteristics of the firm (number of establishments, industry, geographical location); on the level and composition of employment (employment stock and hiring by gender, type of contract, and level of skill); on earnings (average nominal monthly wages and bonuses by worker qualification); working hours (contractual working hours, overtime work, hours lost, number of workers involved in overnight shifts); union status (union representation, number of union in the firms, company wage bargaining).

[11] We concentrated on one-plant firms because for some crucial variables such as the wage we only have information at the firm but not at the plant level.

we chose to deal only with interior solutions and we therefore kept in the sample only firms choosing to hire both open-ended and fixed-term workers. We needed to impose this additional restriction because we do not have the required information to model the two-corner solutions (not hiring on fixed-term contract or not hiring on open-ended contract) without resorting to strong functional assumptions. Therefore our results are conditional on firms that have hired with both types of contracts. We end up with about 307 firms. Table 1, column 1, includes some basic information about the most important variables used in the empirical analysis. On average, selected firms are of medium size with about 200 employees; nominal monthly wages are about € 1600, with those paid to fixed-term employees being about 30 per cent lower than those of open-ended workers.[12] On average our firms hire about 26 workers every year, which corresponds to an inflow rate of about 13 per cent; about 61 per cent of all newly hired employees are hired on a fixed-term contract. About 25 per cent of the employees are female and about 68 per cent are blue-collar workers. Workers holding a fixed-term contract represent about 11 per cent of the stock of employees. In order to evaluate how different our sample is from the rest of the firms included in the Federmeccanica survey, in column 3 of Table 1 we report the mean values computed on the sample enlarged to include those firms that do not meet the criteria of being present in both 2001 and 2002. The most important difference is that the average size shrinks to about 150 employees. However, the structural characteristic of the two samples remains about the same. The average wage is € 1560 (about 3 per cent less than our sample), the share of new workers to total employment is about 14 per cent (one percentage point more), and its composition in term of fixed-term versus open-ended contracts is virtually unchanged. Including in the sample firms that have not hired new workers on fixed-term and open-ended contracts does not alter the qualitative characteristics of the sample except

[12] This ratio represents approximately the actual wage gap between open-ended and fixed-term workers since we do not have information on individual workers' wages, nor on the average wage by type of contract. We only have average wages for each of the 16 job types corresponding to each different contractual position (we do not consider apprenticeship, which is a special contract targeted to young workers) defined by the collective agreement for the industry. These positions are grouped into three main categories (blue-collars, intermediate positions and white-collars) and within each group they are ranked from the least to the most skilled. As a proxy for the wage of fixed-term contracts we use the average wage of the two least skilled groups among the blue-collars and the least skilled workers among the white-collars. We use the remaining categories to compute the wages for open-ended workers. Empirical evidence confirms that these assumptions are not unreasonable. Indeed, according to our data-set, in the period 1998-2002 about 73 out of 100 blue-collar workers were hired on a fixed-term contract, while the corresponding share among white-collars was only 31 per cent. About 87 per cent of all newly hired fixed-term workers were blue-collars, whose share of total hires (both open-ended and fixed-term) was about 74 per cent. Using these assumptions on wage measures, we found that in our sample fixed-term workers earn about 30 per cent less than open-ended ones, a number that is not far from the differential estimated using data from the Bank of Italy Survey on Household Income and Wealth. Using these

for the average size of the firms, which shrinks to about 90 employees. We conclude that our selection criteria do not alter dramatically the characteristics of the Federmeccanica survey except for the average size of the firms.

A preliminary look at the log of new workers hired on fixed-term compared with open-ended contracts before and after the tax credit reveals that something occurred in 2001 and 2002. In Figure 1 we have plotted, for every year in our sample, the empirical distribution of firms with respect to the hiring composition by type of contract, distinguishing between those that increased the stock of open-ended workers and those that did not. We used this criterion because it is the rule that allowed firms to receive the labour cost cut when the tax credit came to force in 2001. The first panel compares the cumulative distribution of raw data. Although firms that increase the share of open-ended workers over the stock of the previous year tend systematically to hire fewer fixed-term than open-ended workers, the difference becomes larger in both 2001 and 2002 compared with previous periods. The change in the composition of hiring appears even more noticeable when we control for firms' fixed effect. Before the tax credit the two cumulative distributions were almost one on top of the other; they spread apart in 2001 and especially in 2002. However, to evaluate how strong this shift is we turn to the econometric exercise.

5. Identification assumptions

Following the analysis of Section 2, the empirical counterparts of the two reduced forms (3) and (4) can be written as

$$(8)\log\left(\frac{H^{ft}}{H^{oe}}\right)_{it} = \tilde{b}_0 + \tilde{b}_{1t}DTC_{it} + \tilde{b}_2 Z_{it} + \tilde{v}_i + \tilde{\varepsilon}_t + \tilde{\eta}_{it}$$

$$(9)\log\left(\frac{W^{ft}}{W^{oe}}\right)_{it} = \tilde{c}_0 + \tilde{c}_{1t}DTC_{it} + \tilde{c}_2 Z_{it} + \tilde{\phi}_i + \tilde{\varepsilon}_t + \tilde{\mu}_{it},$$

with $i = 1.....307, t = 1998, 1999, 2001, 2002$

Here we do not use the direct measure of the labour cost reduction induced by the tax credit (TC_{it}) but rather DTC_{it}, i.e. a dummy variable indicating treatment status in year t; it takes

data for the year 2000, Cipollone and Guelfi (2003) estimate a 32 per cent raw differential for males and a 9 per cent one for females.

value 1 if firm *i* has increased the stock of permanent employees in year *t* and zero otherwise[13]. As the tax credit was started in 2001 we expect $\widetilde{b}_{1t}$ and $\widetilde{c}_{1t}$ to be zero for t<2001. This is the easiest way to identify the effect of the tax credit without resorting to wage information. The rationale for choosing the dummies rather than the actual value for the tax credit is that, as discussed, we do not have a precise measure for the wages paid to workers with different types of contracts. While our approximation seems to be reasonable, it can still weaken the robustness of our results. To overcome this problem we focus only on those results that can be obtained without using any measure of relative wages. Equation (8) allows us to identify the important parameter $\widetilde{b}_{1t}$ that can still be used to evaluate the *VFC* by computing the change in the share of open-ended contracts in total recruitments as $\Delta\left(\frac{H^{oe}}{H^{ft}+H^{oe}}\right)=\frac{1}{k_{it}\exp(\widetilde{b}_{1t})+1}-\frac{1}{k_{it}+1}$.

There are three major econometric difficulties in directly estimating equation (8). The first problem concerns the nature of the effect that we can actually identify. Equation (8) implicitly assumes that the effect of the tax credit on the composition of recruitment is the same for each firm, if different over time. However, if there are heterogeneous reactions to the tax credit, the only parameter that is identified under the standard assumptions (discussed below) is the average Average Treatment on the Treated (ATT), that is the effect of the tax credit for those firms that actually used it. Nothing can be said about the effect of the tax credit for the average firm in the sample.

The other two problems concern the actual possibility of recovering the ATT in our specific context. Identification of $\widetilde{b}_{1t}$ impinges on the fact that the tax credit is not correlated to unobserved shocks to the firm's composition of hiring by type of contract. One general concern regards time-varying firm-specific effects, since those that are constant over time are controlled for by firms' fixed effects. Identification problems could arise if firms that in 2001-2002 increased the share of open-ended workers were reacting to some specific shock rather than to the tax credit. In this instance, the OLS estimates of the slopes $\widetilde{b}_{1t}$ measure the correlations between this shock and the tax credit rather than the causal effect of the labour cost cut on relative labour demand. On top of this standard problem we face the additional

[13] As stated in Section 3 firms received the tax credit if they increased the share of open-ended worker in total employment compared with the same share in the preceding year and the worker hired on permanent job was at least 25 year old. Unfortunately our data do not allow us to control for this last condition. Therefore we

difficulty that the selection rule into treatment – firms are entitled to the tax credit if they increase the number of workers hired on open-ended contracts – might induce a mechanical correlation between the indicator for the treated status DTC_{it}, and the outcome variable (the log of the ratio between fixed-term and open-ended contracts).

To discuss the identification assumptions we need in the face of these two problems let us simplify equation (8) and assume that Z variables are not there:

$$(8') \; \log\left(\frac{H^{ft}}{H^{oe}}\right)_{it} = b_o + \tilde{b}_t DTC_{it} + \zeta_{it}$$

OLS on 8' return

$$(10) \; \tilde{b}_{1tOLS} = \tilde{b}_{1t} + \frac{\mathrm{cov}(DTC_{it}, \zeta_{it})}{Var(DTC_{it})}$$

for t=2001 or t=2002, as in these two years the tax credit was in force; in contrast for t=1998 or t=1999 OLS on 8' return,

$$(10') \; \tilde{b}_{1tOLS} = \frac{\mathrm{cov}(DTC_{it}, \zeta_{it})}{Var(DTC_{it})}$$

as in these years there could be no causal effect of the tax credit on the composition of the new workers by type of contract.

First note that if $\tilde{b}_{1tOLS} = 0$ for t< 2000 then OLS coefficients of DTC_{it} for t> 2000 identify the causal effect of the tax credit on the log of composition of the recruits ($\tilde{b}_{1tOLS} = \tilde{b}_{1t}$ for t> 2000), provided that the covariance between the error term and the indicator of treatment status is constant over time, that is $\frac{\mathrm{cov}(DTC_{it}, \zeta_{it})}{Var(DTC_{it})} = \frac{\mathrm{cov}(DTC_{it'}, \zeta_{it'})}{Var(DTC_{it'})}$. This is the standard assumption for a diff-in-diff estimator. In the case in which $\tilde{b}_{1tOLS} \neq 0$ for t< 2000 we can still recover the true effect of the tax credit by taking the difference between *b*'s estimate for a *t* in which the tax credit was in force and a *t'* without tax credit

$$(11) \; ATT_t = \tilde{b}_{1tOLS} - \tilde{b}_{1t'OLS} = \tilde{b}_{1t} + \frac{\mathrm{cov}(DTC_{it}, \zeta_{it})}{Var(DTC_{it})} - \frac{\mathrm{cov}(DTC_{it'}, \zeta_{it'})}{Var(DTC_{it'})}$$

As in the case in which $\tilde{b}_{1tOLS} = 0$ for t<2000, this difference recovers the true effect as long as the bias due to the covariance between treated status and error term is constant

measure with error the treated status of firms and the estimated effect of the tax credit can be regarded as a lower bound of the actual effect.

over time. To be more explicit about the assumption needed to identify the true effect of the tax credit, regardless of the value of $\widetilde{b}_{1tOLS}$ for t< 2000, let us assume that the error term has three components:

$$(12)\ \zeta_{it} = \widetilde{v}_{it} + (\sigma^{T}\varepsilon_{t} + \eta_{it}^{T})DTC_{it} + (\sigma^{C}\varepsilon_{t} + \eta_{it}^{C})(1 - DTC_{it})$$

with both time trend (ε_t) and time-varying firm-specific shock (η_{it}) differing between firms that increase permanent employment (superscript T) and those that do not (superscript C). With this structure of the error terms and recalling that DTC_{it} is a dummy variable the bias term in 11 can be written as:

$$\frac{\mathrm{cov}(DTC_{it},\zeta_{it})}{Var(DTC_{it})} - \frac{\mathrm{cov}(DTC_{it'},\zeta_{it'})}{Var(DTC_{it'})} = A1 + A2 + A3$$

where

$A1 = \{E[\widetilde{v}_i \mid DTC_{it} = 1] - E[\widetilde{v}_i \mid DTC_{it} = 0]\} - \{E[\widetilde{v}_i \mid DTC_{it'} = 1] - E[\widetilde{v}_i \mid DTC_{it'} = 0]\}$ is due to the difference in the composition of the treated and control group over time; $A2 = (\sigma^{T} - \sigma^{C})(\varepsilon_t - \varepsilon_{t'})$ is due to the difference in the trend between treated and control group, and $A3 = \{E[\eta_{it} \mid DTC_{it} = 1] - E[\eta_{it} \mid DTC_{it} = 0]\} - \{E[\eta_{it'} \mid DTC_{it'} = 1] - E[\eta_{it'} \mid DTC_{it'} = 0]\}$ is the component due to the potential self-selection of firms into the treatment because of differential firm time-specific shock in the years when the tax credit was in force or due to the mechanical correlation between the indicator of treatment status and the year firm-specific shock. In a standard diff-in-diff framework one needs to assume that these three components are all zero.

However, our data sets allows us to relax some of these assumptions. Since we have a panel of firms the bias component due to specific fixed effect can be controlled for by means of a fixed effect estimator. In order to evaluate the importance of the second and third component of the bias we can compare two $\widetilde{b}_{1tOLS}$ estimated for two years in which the tax credit was not in force, such as 1999 and 1998.[14] If these coefficients are all zero then the bias in equation (10) is zero and we can identify the effect of the tax credit as a simple difference between the mean outcome of treated and control units; if they are different from zero but their difference is zero then the standard assumption for a diff-in-diff holds and we can identify the effect of the tax credit as the difference between two betas computed for

[14] This is basically the same idea as that underlying the Adjusted for Differential Trend Diff-in-Diff estimator proposed by Bell et al. (1999).

years with and without the tax credit in force. This case would emerge if, for example, the joint distribution of the specific time-firm-specific shock (η_{it}) and the indicator of the treatment status (DCT_{it}) did not depend on time and if the trend were common to both the treated and the control group ($\sigma^T - \sigma^C = 0$). Finally, a non-zero difference might suggest that the A2 and A3 components of the bias affect the estimate of the effect of the tax credit. We can take into account this potential distortion using as estimator of the causal effect of the tax credit the difference of two differences among betas, one computed on years with and without the tax credit in force and one comparing two years without the tax credit; in our specific context we can implement this strategy as: $(\tilde{b}_{1,2002OLS} - \tilde{b}_{1,1999OLS}) - (\tilde{b}_{1,1999OLS} - \tilde{b}_{1,1998OLS})$.

The weakness of the parametric approach of equation (8') is that it attributes the same weight to all the control units, regardless of their closeness to the treated units. This characteristic might be a problem if the observable variables that drive the selection process into the treatment status do not have the same support for treated and control units, or have different distribution over the common support. To take into account these additional problems we rely on the matching estimators based on the probability of being treated (Heckman et al. 1997, Heckman et alt. 1998 and Blundel and Costa Diaz 2000). For each year of our sample we compute a matching estimator based on the propensity score:

$$(12)\ ATT_{Mt} = \sum_{i \in \{DCT_{it}=1\}} \left[\log(\frac{H_{it}^{ft}}{H_{it}^{oe}}) - \sum_{j \in \{DCT_{jt}=0\}} W_{ijt} \log(\frac{H_{jt}^{ft}}{H_{jt}^{oe}}) \right] w_{it}$$

where w_{it} is the weight of the *i-th* treated observation at time *t* and W_{ijt} represents the weight that it is given when comparing treated observation *i* with control observation *j* at time *t*. There are many ways of implementing this matching estimator depending on the algorithm used to compute W_{ijt}. We use the Kernel matching version implemented by Becker and Ichino (2002)

$$(13)\ ATT_{KMt} = \left(\frac{1}{\sum_i DCT_{it}} \right) \sum_{i \in \{DCT_{it}=1\}} \left\{ \log(\frac{H_{it}^{ft}}{H_{it}^{oe}}) - \sum_{j \in \{DCT_{jt}=0\}} \left[\frac{G(\frac{p_j - p_i}{h_n}) \log(\frac{H_{jt}^{ft}}{H_{jt}^{oe}})}{\sum_{j \in \{DCT_{ij}=0\}} G(\frac{p_j - p_i}{h_n})} \right] \right\}$$

where G(·) is a kernel function, p_j and p_i are the estimated propensity scores – i.e. the probability of being treated – of observations *j* and *i* at time *t*, and h_n is the bandwidth.

In analogy with our discussion of the parametric case, we can directly read the causal effect of the tax credit from the estimator ATT_{KMt} for t> 2000 if ATT_{KMt} =0 for t< 2002; if this condition does not hold we can still recover the causal effect of the tax credit by computing the difference between two matching estimators estimated one in a year with the tax credit and one without.

6. Results

In order to implement the strategy described in the previous section we begin by estimating equation (8) with OLS by pooling all the observations for the period 1998, 1999, 2001 and 2002

$$(14)\log\left(\frac{H^{ft}}{H^{oe}}\right)_{it} = b_0 + b_2 Z_{it} + \sum_t \widetilde{b}_{1,t} DTC_{it} + \sum_t \widetilde{\varepsilon}_t + \widetilde{v}_i + \widetilde{\eta}_{it}$$

In the second column of Table 2 we present the estimated $\widetilde{b}_{1_t}$ in a model without any type of control (not even fixed effect) beside the year dummies. The betas for the years in which the tax credit was in force, namely 2001 and 2002, are in the order of -.40 with a standard error in the order of .15. The betas for the years in which the tax credit was not in force are in the order of -.30 with a standard error of .20. Reading these results at their face values one would conclude that the betas in the years in which the tax credit was in force identify the true causal effect of the tax credit on the composition of recruitments as the covariance between the indicator of the treated status and the error term is mostly noise. As discussed in the previous section, the implication of this finding is that even a simple difference between the outcome of treated and control units in the year for which the tax credit was in force should be able to recover the casual effect of the tax credit on the composition of the hirings. Therefore one would conclude that this causal effect is about -.40. However, the point estimate of the treatment status for the year before 2000 is large enough to make us uncomfortable about discarding its implication on the ground that it not significantly different from zero. This consideration leads us to include more controls in the equation. In column 3 of Table 2 we estimate the equation including the usual treatment status indicator along with the firm fixed effect. Adjusting for time-invariant characteristics of the firm it further increases the slope of the indicator of the treated status for the years in which the tax credit was in force (from -.40 to about -.45 on the average for 2002 and 2001), while at the same time it shrinks the same slopes for the years in which the tax credit was not there (from -.32 to -.25 in 1999 and from -.30 to -.11). The precision of these last two estimates improves marginally but not enough to take their statistical significance above the usual threshold. Hence we cannot reject the hypothesis that the bias in equation 10 is equal to zero. On the basis of these findings we restate that the effect of the tax credit on the composition of hirings is of the order of -.45.

Column 4 of Table 2 presents the estimate of model 14 with a series of additional controls. We include controls for the gender and skill composition of employees and for the share of workers involved in overnight shifts to account for possible firm-specific trends in the composition of the labour force. For instance, such an occurrence would be brought about by the need to accommodate technical upgrading of the production process, or to meet quality standards required by new customers.[15] In these cases the share of open-ended workers increases for reasons other than the tax credit. Firms' fixed effect cannot capture this trend. Moreover, as the composition by type of contract of the flow of new workers might reflect the composition of pre-existing stocks, we also include in the equation the share of fixed-term workers relative to total employment in the previous year. Finally, as the composition of recruitments might be influenced by firms' specific business cycle we include in the equation the (log) number of hours of overtime work, hours of wage supplementation funds and hours lost because of strikes.

All these additional controls reinforce the pattern that emerged in the previous two models. The coefficients of the treated status increase in 2001 and especially in 2002, rising on average to -.50, but remain constant in 1999 and drop further in 1998. As the standard errors seem rather unaffected by the inclusion of these additional controls the statistical significance of the last two coefficients weakens further.

We conclude that the tax credit seems to have had a causal effect on the composition of hirings by reducing the share of fixed-term contracts in favour of open-ended contracts. Using the most conservative of the above results we can quantify this effect at -.39, which implies that the share of workers hired on open-ended contracts out of all recruits increased by 7.6 percentage points because of the tax credit. As the average number of recruits in our sample equals 26, our result implies that the tax credit induced firms to hire two workers on open-ended contracts that would have been hired on fixed-term contracts otherwise.

As discussed in the previous section, direct comparison of the average outcome for the treated and the control group could produce highly biased results if the two groups do not

[15] These examples can be expressed as an increase in the ω_{it} parameter of the production function (5).

share the same support of the observable variables or have different distributions of these variables within a common support.

To address this problem we have implemented the matching estimator (13). As a preliminary step we have estimated the propensity score (the probability of each observation of being treated[16]) and plotted in Figure 2 a panel for every year. Predicted probabilities are reported on the vertical axis; the horizontal axis presents the actual treated status with a value of 0 meaning "not treated" and a value of 1 "treated". Each box presents the distribution of firms by predicted probability of being treated. It begins at the 25th percentile and ends at the 75th percentile; the line in the middle represents the median of the distribution.[17] This graph allows us to evaluate whether treated and control units have propensity scores located on a common support. In our case about 50 per cent of the observations are estimated to have a probability of being treated in the range of .55-.65. The lack of an appreciable difference in the predicted probabilities of being treated between treated and control units suggests that the observed variables used to estimate the propensity score account for most of the selection in the treatment. Note also that the estimated probabilities change little over time and keep a good degree of overlapping between treated and control group. These results allow us to safely compare treated and control, as the members of the two groupa have a close probability of being treated.

Table 3 presents the results of the matching estimator. Column 2 reports the estimator (13) computed on the row data so that it can be regarded as a non-parametric version of model 1 in Table 2. Results are almost identical to those of Table 3. In the years in which the tax credit was in force (2001 and 2002) the difference in the log share of fixed to open-ended contract outcomes between treated and control firms is about -.42 with a standard error in the order of .15.[18] In contrast when the tax credit was not in force the point estimate shrinks and the standard errors increase; as a result the effect of the treatment status in these years cannot be distinguished from zero. Taken at the face value this result would imply that the effect of the tax credit is in the order of -.40. However, the same considerations discussed for the parametric case apply here. Therefore we closely followed the parametric analysis

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[16] We use the implementation of Becker and Ichino (2002) and we balance the observed variables used to estimate the propensity score within each of the five groups in which we divide the distribution of the propensity score.

[17] The line emerging from below the box extends up to the data point equal to pc25-1.5*(pc75-pc25), where pc25 is the 25th percentile and pc75 is the 75th percentile. The line emerging from above the box extends up to the data point equal to pc75-1.5*(pc75-pc25).

[18] Standard errors are computed by bootstrap with 100 replications.

and recomputed the ATT_{KM} estimator (13) on the share of fixed-term to open-ended contracts purged by the firm's fixed effects (column 3) and the other covariates used in Table 2, column 4. The matched estimators applied to data purged of the firm's fixed effects basically tell the same story of a negative affect in the years in which the tax credit was in force, although smaller than before (-.33 instead of -.40), and a smaller and not statistically significant effect in the years without tax credit. This pattern is reinforced when we use data purged by fixed effect and other observable variables. Taking the most conservative of these non-parametric estimates we can conclude that the tax credit induced a shift in the log share of fixed to open-ended contracts of about -.18, which implies that the share of recruits with open-ended on all recruits increases by about 3.2 percentage points; in absolute terms this result implies that, on average for the sample, the tax credit induced firms to shift one new worker from a fixed-term to an open-ended contract.

7. Evaluation of the value of flexible contracts

Estimates of the reaction of the log of relative demand $\log(\frac{H^{ft}}{H^{oe}})$ serve to compute the value of flexible contracts (VFC). Note, however, that the reduction in labour cost due to tax credit (i.e. the numerator of the formula defying the VFC) was limited in time, lasting at most three years. Therefore the expected cut in labour cost associated with the tax credit changes with the expected tenure: the longer the duration of the employment spell of the newly hired permanent worker, the lower the reduction in the labour cost associated with the tax credit. In Graph 3 we compute the value of flexible contracts for an employment spell ranging from 1 to 30 years.[19] The two lines in the graph are computed using as a measure of the effect of the tax credit the most conservative estimates obtained with the parametric method (-.39 in Table 2) and the non-parametric method (-.18 in Table 3).

According to our calculation firms would have valued the opportunity to hire 1 per cent of their new workers on fixed-term contracts rather than open-ended contracts as much as a reduction in the wage of open-ended workers of between 3 and 7 percentage points in the first three years of the employment spell.[20] Thereafter the value of the flexible contracts

[19] To compute the value of flexible contract we have made the following assumption: real wage constant at the 2001 value, real interest rate of 1 per cent and North-South composition calculated on the basis of employees in industry excluding construction. The reference labour cost is the average of the sample

[20] Here we derived the value of the flexible contract by looking only at the empirical counterpart of equation (3). As explained in Section 2 this is only an approximation of the actual value that firms attribute to the flexibility. However, the quality of this approximation depends on the coefficient $\tilde{c}_{1t}$ in equation (9); the

declines when the reduction in the labour cost due to the tax credit shrinks as a proportion of the overall wage bill of the new worker. After 5 years the VFC has already dropped within the range of 1.9-4.4 per cent and to between 0.4 and 0.8 per cent for employment spells lasting 30 years. As in Italy the average duration of employment spells is estimated at around 8 years (Cingano and Rosolia 2003, Garibaldi and Violante 2005), we conclude that the possibility of hiring 1 per cent of new workers with a fixed-term contract is equivalent, on average, to a permanent reduction in the wage of an open-ended worker of anything between 1.3 and 2.8 per cent.

Since the average firm in the sample hires about 26 people a year and therefore one new workers is equivalent to 3.8 per cent of new staff, our result suggests that an average firm would be indifferent between hiring one worker on a fixed-term contract and the same worker on an open-ended contract provided his wage is reduced by anything between 5 and 10.7 percentage points.

8. Implications for aggregate employment growth

The results of the previous section can help to explain the rather puzzling dynamics observed in Italian employment in recent years. Between 1995 and 2003 total employment in the Italian private sector grew by about 1.4 million, i.e. an overall increase of 8.4 per cent. In the same period value added rose by 13.7 per cent. The ex-post elasticity was about 0.6, a very large and unprecedented value by Italian standards; in the previous fifteen years employment declined by 4.5 per cent while output increased by 37 per cent. Wage moderation and the introduction of flexible contracts have been invoked to explain this new feature of the Italian economy (Brandolini at al. 2005). Indeed real wages remained basically constant over the period 1995-2003 and the share of fixed-term contracts in all newly hired employees grew from 34 to 42 per cent. However, the actual importance of the new flexibility for the expansion of employment has yet to be quantified. Our results help to fill the gap. They indicate that firms in the engineering sector did value the growth of fixed-term contracts as equivalent to a cut in the wage of an open-ended worker of anything between 10 and 22 per cent. In Italy the long-run elasticity of employment to wages is estimated at around -0.3; therefore, applying the value of flexible contract for engineering firms to the

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smaller this coefficient the better the approximation. We have estimated this coefficient using the same technique as for the estimation of $\widetilde{b}_{1t}$ and our results suggest that it is close to zero.

whole private sector, we can calculate that the rise in the share of flexible contracts raised the long-run level of employment by 3.1-6.7 percentage points.

9. Conclusions

In this paper we estimate the value firms attach to flexible labour contracts. We rely on a panel of Italian firms operating in the engineering sector in the period 1998-2002. We estimate a demand curve for flexible contracts relative to permanent ones by exploiting a reduction in the labour cost of open-ended jobs granted to firms hiring new workers with open-ended rather than fixed-term contracts. Identification is achieved by comparing how the composition of recruitment by type of contract changed for firms that received the tax credit and those that did not. Potential distortions due to self-selection into the programme, because of firm-specific time-varying shocks or the mechanical correlation induced by the selection rule into the programme, are accounted for by estimating the spurious effect of the tax credit in the years when it was not in force. We found that firms appraised the possibility of hiring 1 per cent of new workers on temporary contracts as much as a reduction in the labour cost of a worker on an open-ended contract in the range of 1.3-2.8 per cent. As the share of fixed-term contracts in all newly hired employees grew by 8 percentage points between 1995 and 2003 (from 34 to 42 per cent), the advent of flexibility in the Italian labour market was worth as much as a drop in the wage of permanent workers in the range of 10-22 per cent. This labour cost cut might explain the large increase in Italian employment (between 3.1 and 6.7 of the overall 8.4 percentage points), which otherwise remains rather puzzling in the light of the slow growth of output.

Tables and figures

Table 1

BASIC SAMPLE CHARACTERISTICS

	Firms in the sample:	For comparison	
	All firms with a single plant and positive recruits of both open-ended and fixed-term and in the sample in 2001 and 2002	All firms with a single plant and positive recruits of both open-ended and fixed-term	All firms with a single plant
Wage(1)	1596	1559	1546
Open-ended	1639	1596	1602
Fixed-term	1130	1091	1098
Recruits	26	22	12
Open-ended	10	8	5
Fixed-term	16	13	7
Employees	205	149	89
Women	40	31	18
Blue collars	139	94	57
Stock of fixed-term workers (2)	0.11	0.12	0.08
Employees involved in shifts (2)	0.66	0.66	0.65
Wage Supplementation Fund (3)	0.01	0.01	0.01
Over time (3)	92	98	81
Strikes (3)	7.4	5.5	4.0
Number of observations	938	2941	9089
Number of firms	307	1685	4599

(1) Euros. (2) As a share of total employment. (3) Hours per employee.

Source: Own computations on Federmeccanica data.

Table 2

ESTIMATES OF THE EFFECT OF THE TAX CREDIT ON RELATIVE LABOUR DEMAND FOR FIXED-TERM WORKERS. LINEAR MODELS

(Dependent variable: log of the share of fixed-term new recruits relative to open-ended [1])

	Model 1: no adjustment	**Model 2: Adjusted for firms' fixed effects**	**Model 3: Adjusted for firms' fixed effects and other covariates[2]**
Treated in 2002 ($\widetilde{b}_{1,2002}$)	**-0.41**	**-0.44**	**-0.54**
	0.14	*0.13*	*0.14*
Treated in 2001 ($\widetilde{b}_{1,2001}$)	**-0.39**	**-0.45**	**-0.47**
	0.15	*0.14*	*0.15*
Treated in 1999 ($\widetilde{b}_{1,1999}$)	**-0.32**	**-0.25**	**-0.25**
	0.20	*0.17*	*0.17*
Treated in 1998 ($\widetilde{b}_{1,1998}$)	**-0.30**	**-0.11**	**-0.08**
	0.20	*0.17*	*0.18*
Years' fixed effect	*Yes*	*Yes*	*Yes*
Firms' fixed effect		*Yes*	*Yes*
Other controls			*Yes*
Number of firms	*307*	*307*	*307*
Number of observations	*938*	*938*	*938*

[1] Estimate of equation 12 in the text for the years 1998, 1999, 2001, 2002. Standard errors in italics. [2] Number of employees, share of women, blue-collar workers, open-ended workers at time *t-1,* workers involved in overnight shifts, log of per worker hours of overtime , of Wage Supplementation Fund and strikes.

Source: Own computations on Federmeccanica data

Table 3

ESTIMATES OF THE EFFECT OF THE TAX CREDIT ON RELATIVE LABOUR DEMAND FOR FIXED-TERM WORKERS. KERNEL MATCHING ESTIMATE

(Dependent variable: log of the share of fixed-term new recruits relative to open-ended [1])

	Model 1: no adjustment	**Model 2: Adjusted for firms' fixed effects**	**Model 3: Adjusted for firms' fixed effects and other covariates[2]**
ATT_{KM} in 2002	**-0.45**	**-0.39**	**-0.27**
	0.19	*0.10*	*0.09*
ATT_{KM} in 2001	**-0.39**	**-0.27**	**-0.18**
	0.15	*0.10*	*0.08*
ATT_{KM} in 1999	**-0.31**	**-0.18**	**-0.12**
	0.21	*0.11*	*0.13*
ATT_{KM} in 1998	**-0.27**	**-0.11**	**0.01**
	0.20	*0.15*	*0.13*
Number of firms	*307*	*307*	*307*
Number of observations	*938*	*938*	*938*

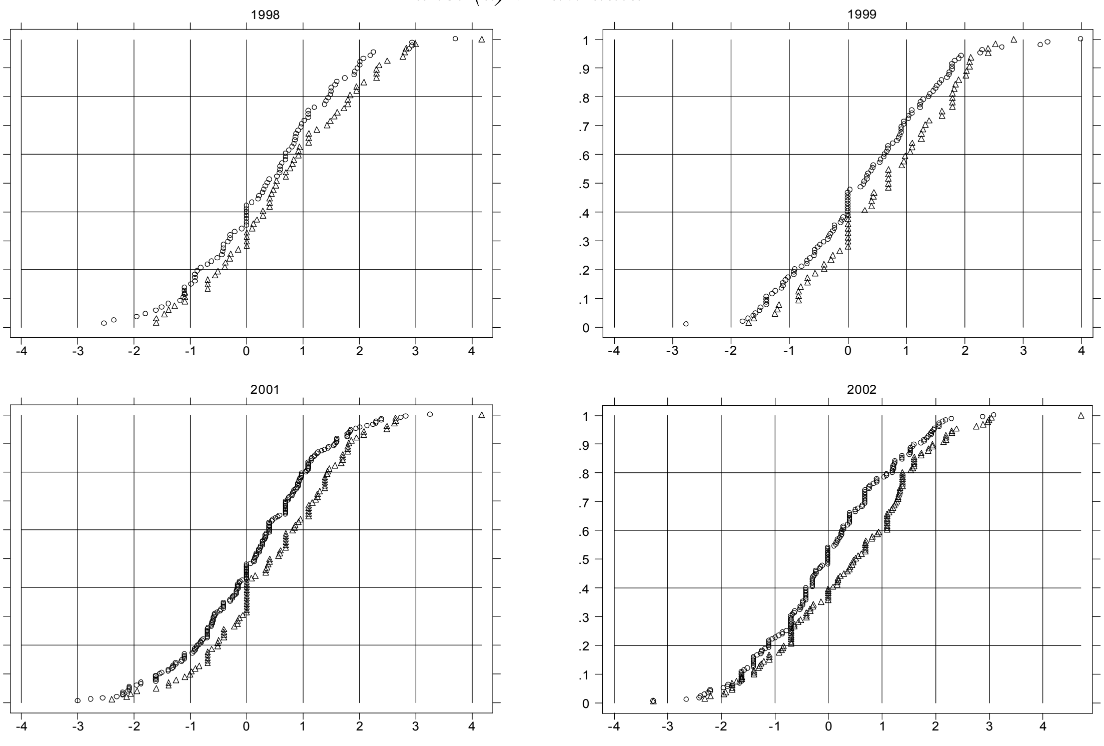
[1] Estimator presented in equation 13 in the text for the years 1998, 1999, 2001, and 2002. Standard errors in italics, they are computed by bootstrap with 100 repetitions. [2] Number of employees, share of women, blue-collar workers, open-ended workers at time *t-1,* workers involved in overnight shifts, log of per worker hours of overtime, of Wage Supplementation Fund and strikes.

Source: Own computations on Federmeccanica data

Figure 1

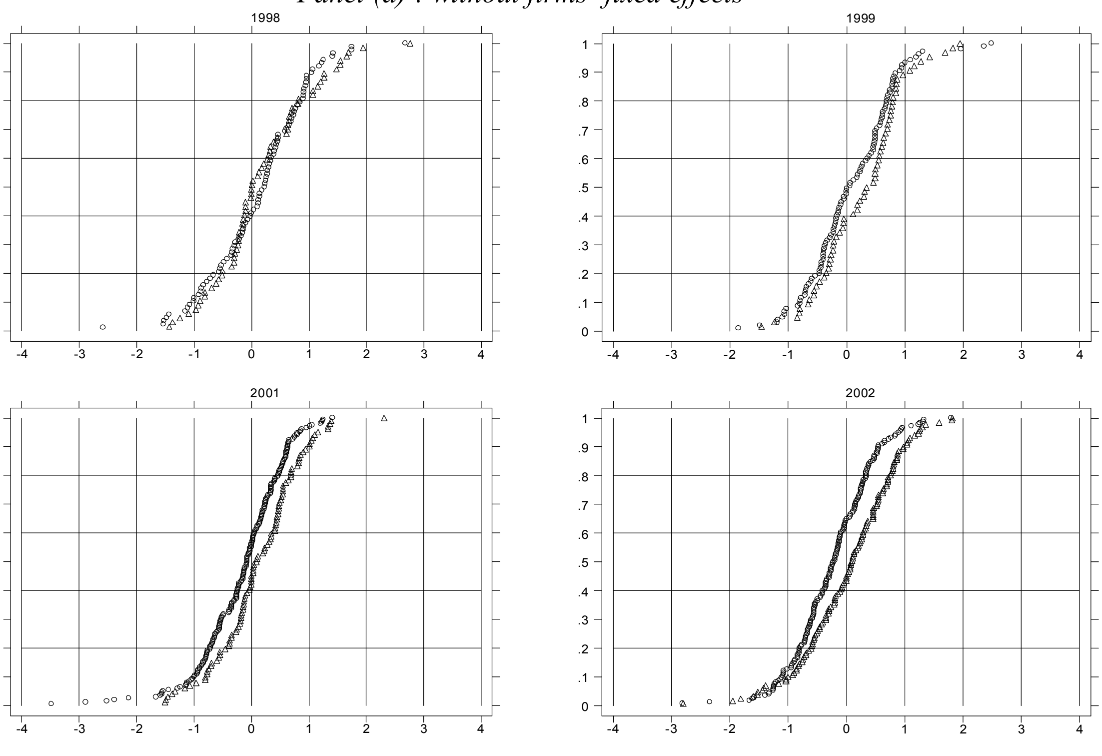
CUMULATVE DENSITIES OF THE LOG OF FIXED-TERM RELATIVE TO OPEN-ENDED HIRES

Panel (a) : Raw data



Log of the ratio of recruits with fixed-term to recruits with open-ended contracts

Panel (a) : without firms' fixed effects



Log of the ratio of recruits with fixed-term to recruits with open-ended contracts

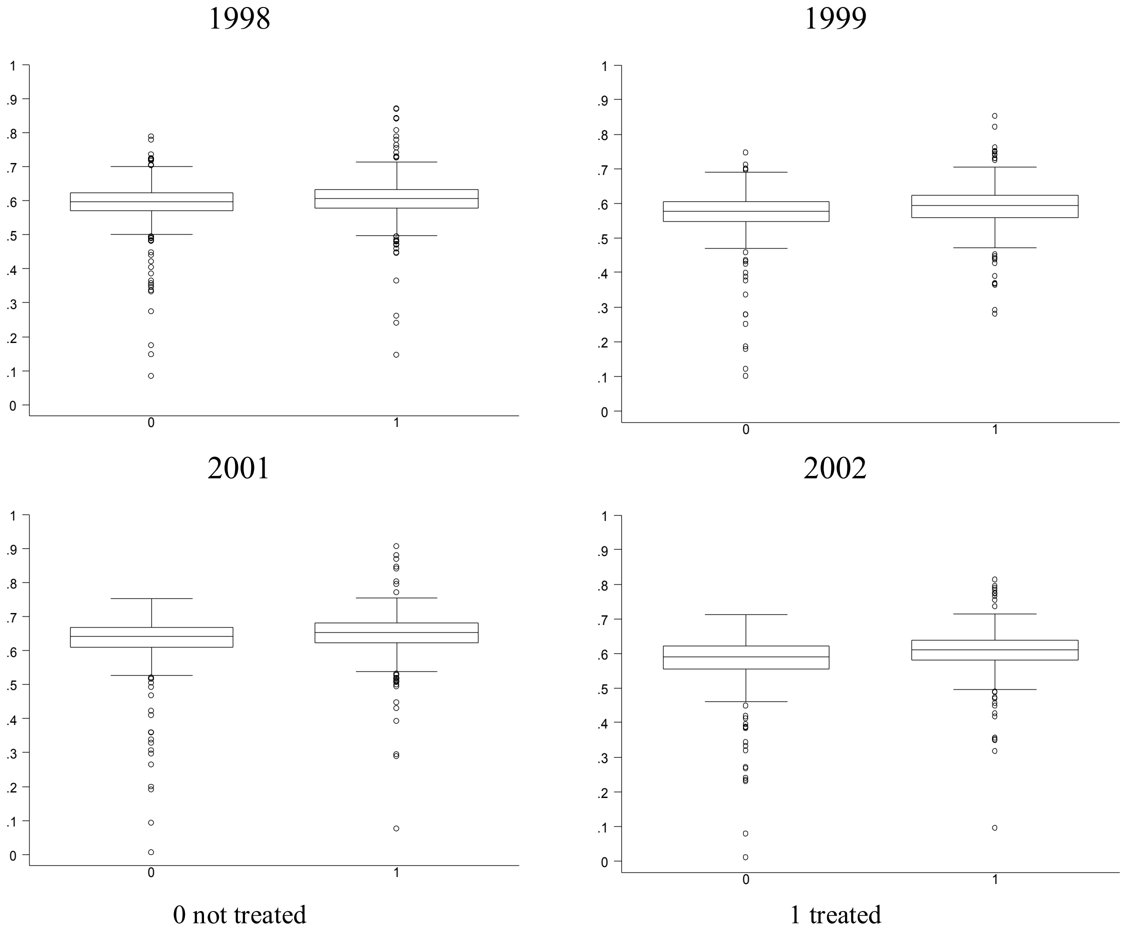
○ Firms increasing the stock of open-ended workers

∆ Firms not increasing the stock of open-ended workers

Source: Own computations on Federmeccanica data.

Figure 2

PREDICTED PROBABILITY OF BEING TREATED BY YEAR AND TREATED STATUS[1]

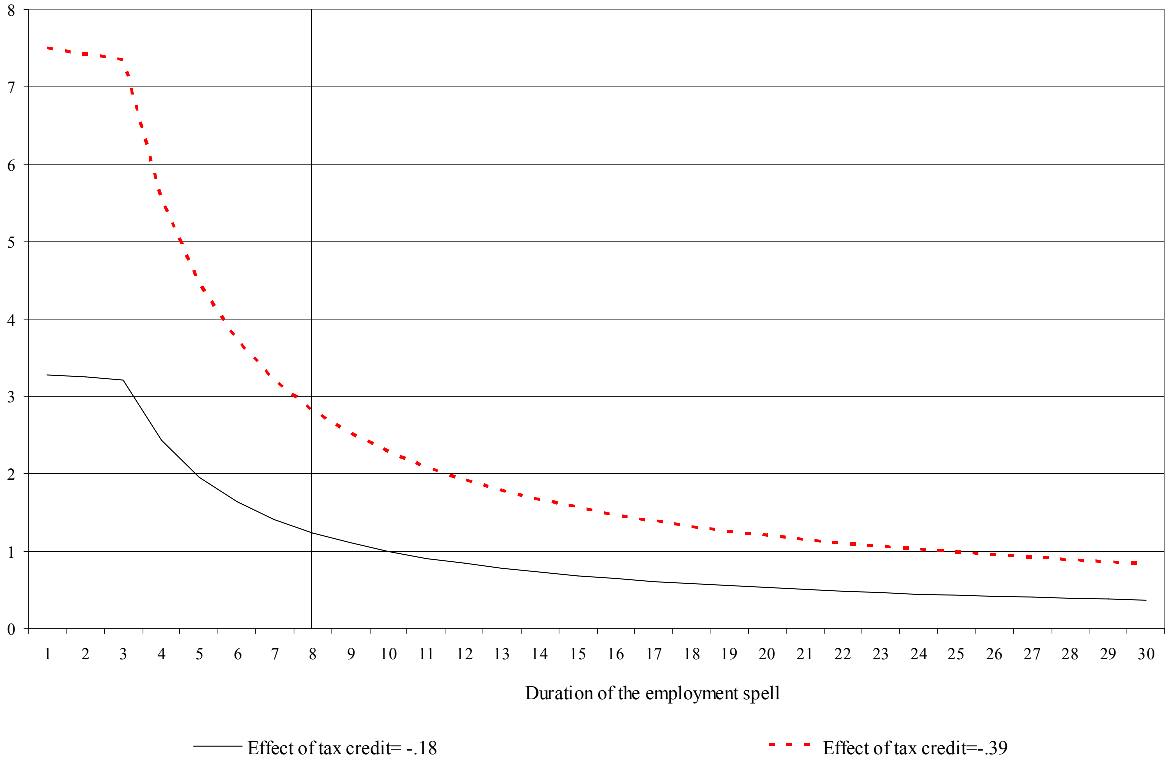


[1] The probit model includes as regressors the log of the number of employees of the firm, the share of women, blue-collar workers, open-ended workers at time *t-1,* share of workers involved in overnight shifts, log of per worker hours of overtime, of Wage Supplementation Fund, and strikes.

Source: Own computations on Federmeccanica data.

Figure 3

ESTIMATED VALUE OF THE FLEXIBLE CONTRACT BY DURATION OF EMPLYMENT SPELL[1]



[1] The two lines are computed using the formula

$$VFC(T) = \frac{\sum_{t=1}^{T} \lambda^{t-1} VTC_t}{\sum_{t=1}^{T} \lambda^{t-1} w_t} \left[\frac{1}{k \ \exp(\tilde{b}_1) + 1} - \frac{1}{k \ + 1} \right]^{-1}$$

and using $\tilde{b}_1 = -.39$ (dashed line) or $\tilde{b}_1 = -.18$ (solid line); the other parameters are $\lambda_1 = .99; k = 3.52$; labour costs are kept constant at their 2001 values and VTC amount to about € 7500 and € 5000, in the South and in the Centre-North respectively for the first three years, and zero thereafter.

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