

WORKING PAPERS

**Temporary employment,
demand volatility and unions:
Firm-level evidence**

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Temporary employment, demand volatility and unions: Firm-level evidence

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Abstract

This paper investigates the effect of workplace unionization and demand volatility on firms' propensity to use temporary employment. Using Italian firm level data, we show that the impact of unionization crucially depends on the type of fixed term contracts considered and the degree of volatility. The presence of unions *per se* is found to increase the demand for non-training temporary contracts, while it does not affect the demand for training temporary contracts. We argue that this occurs because non-training temporary contracts are typically used as a buffer stock to cope with uncertainty, and the unions tend to encourage their use to protect insider workers. Training temporary contracts, on the other hand, do not lend themselves to this function, as they are more likely to be used as a screening device for future permanent positions. We also find that, as volatility increases, the unions become more concerned about the weakening of their bargaining power and tend to discourage the hiring of non-training temporary workers.

Keywords: unions, temporary workers, training, product demand volatility, firms.

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1. Introduction

In the last few decades, most European countries have set about a series of labor market reforms designed to provide firms with enhanced contractual flexibility in the face of the increasing economic uncertainty related to globalization and technological change. These reforms have generally been introduced “at the margin”, facilitating the use of fixed term contracts for new hires while leaving employment protection for open-ended contracts unaffected. As a result, the diffusion of temporary employment has become one of the distinctive features of the European labor market, ultimately producing dual labor markets with many workers hired on fixed-term contracts entrapped in “lower quality” jobs (Boeri, 2012).¹

For this reason, economists and policy makers are showing increasing interest in identifying the determinants of temporary employment. Most of the literature has generally focused on labor supply dynamics, typically using individual data from labor force surveys (Booth et al, 2002, among others). The firms' demand for temporary workers has been studied rather less despite its considerable significance since the available empirical evidence shows that fixed-term contracts are typically favored by firms much more than by workers (e.g., Francesconi and Garcia-Serrano, 2004).

The institutions and the characteristics of the economic environment in which firms operate clearly constitute important determinants of their demand for temporary employment. The literature has particularly focused on two factors. The first concerns the role of the unions, which are of primary relevance in understanding the employment dynamics in European labor markets. An under-explored issue, which lies at the heart of this paper, is whether the pervasive role of unions and collective bargaining in most EU countries may have contributed to the diffusion of temporary employment, as a way to maximize the utility of the types of workers (“insiders”) that the union most represents. One of the insights emerging from the insider-outsider literature is that

¹ Temporary contracts account for 14% of EU employment and for more than 50% of new hires (OCSE, 2008).

the unions may contribute to labor market duality by favoring the use of temporary employment as a “buffer” to isolate permanent workers from the negative effects of demand uncertainty and technological shocks (Bentolila and Dolado, 1994; Saint Paul, 1996). Moreover, the presence of unions in the workplace may increase the bargaining power of permanent workers (insiders) on wages, thus inducing firms to make use of fixed term contracts to exploit numerical and wage flexibility (Abraham and Taylor, 1996). On the other hand, the unions may also oppose the diffusion of fixed-term contracts, as temporary workers are typically less unionized and less engaged in industrial relations, representing a potential threat to union strength and representativeness (Heery, 2004; Salvatori, 2012). Unions may also discourage an excessive use of fixed-term contracts in consideration of the negative effects on social welfare and labor market cohesion (Visser, 2002).

The second factor that has drawn attention in the literature on the determinants of temporary employment demand relates to the volatility of the economic environment in which firms operate. Abraham and Taylor (1996) and Bentolila and Saint Paul, (1994), among others, argue that in more volatile environments firms increase the use of temporary employment, which may represent a valuable buffer stock to adjust to fluctuations. However, the riskiness of the environment does not only affect the behavior of the firms: the unions’ behavior is also likely to be affected, though this issue has been somewhat neglected by the literature on temporary employment.

To the best of our knowledge this is the first paper that focuses explicitly on the interplay between workplace unionization and volatility in the economic environment in generating demand for temporary employment.

We focus our empirical analysis on Italy, an interesting case to study since, on the one hand, there has been a notable increase in temporary contracts in the last decades and, on the other, the unions have considerable power in bargaining with firms over employment and wages.

We rely on a rich firm-level dataset, the *Employer and Employee Survey (RIL)*, conducted by the *Institute for the Development of Workers Professional Training (ISFOL)* in 2005 and 2007. The data refer to firms operating in the Italian extra-agricultural sector, and provide extensive information about different typologies of temporary contracts used by firms, union presence and other workplace characteristics. In particular, the data enable us to study the effect of unionization on the *share* of the different types of temporary employment, rather than simply on the probability of hiring on temporary contracts (e.g., Salvadori, 2009). We also supplement the RIL data with balance-sheet information from the database AIDA by Bureau van Dijk, which contains the universe of incorporated firms in the non-financial private sector. We use AIDA to construct measures of the economic volatility faced by firms and unions based on past sales at the sectoral level, and to obtain instrumental variables for our econometric analyses.

Our preferred econometric specification regresses the share of temporary workers on union status, volatility, and their interaction, after controlling for an ample set of firm covariates and accounting for potential endogeneity issues.

The main findings of the paper can be summarized as follows. First, we find that the effects of unions and volatility crucially depend on the type of temporary contracts. In particular, the presence of unions *per se* increases the demand for non-training temporary contracts, while its impact is not statistically significant for training temporary contracts. We argue that this result is due to the different nature of the two contractual arrangements. In particular, non-training temporary contracts are typically used by firms, and encouraged by unions, as a buffer stock for insider workers. Training temporary contracts, on the other hand, are more likely to play a screening role for a firm's core-staff needs, and are generally more expensive; hence they cannot represent an equally valid buffer for the union's insiders.

Second, we find that, when a firm faces marked volatility in product demand, unions become more concerned about the weakening of their bargaining power due to the use of temporary

workers, and hence tend to discourage their incidence. No such effects are detected for training contracts.

Overall, the paper shows that, in investigating the effect of unions on the demand for temporary contracts, the literature might hitherto have overlooked the fact that such an effect crucially depends on the type of fixed-term contracts considered, as well as the interplay between the unions' objective function and the general demand conditions faced by the firm.

The paper is organized as follows. In section 2 we review the related literature; section 3 deals with the institutional setting, while Section 4 introduces the data and descriptive statistics. Econometric analysis is performed in Section 5 and Section 6. Section 7 concludes.

2. The related literature

Temporary contracts have been widely studied in the last few decades from different standpoints. In the literature attention has focused mainly on the alleged beneficial impacts in terms of unemployment reduction, especially for young workers (OECD; 2008), as well as the potential negative effects on the welfare of workers (Booth et al., 2002; Blanchard and Landier, 2002; Boeri, 2012) and on labor productivity (Battisti and Vallanti, 2013; Cappellari et al., 2012; Damiani et al., 2011) – effects that are often attributed to the two-tier nature of the institutional reforms.

It is worth noting that the majority of the applied research on the topic is based on individual data that are unlikely to account appropriately for the firms' incentives to hire temporary employment (Booth and Francesconi, 2003). Indeed, the available evidence shows that a substantial part of temporary employment appears to be involuntary for the worker. This suggests that firms' characteristics and personnel policies play a key role in determining the diffusion of fixed-term contracts. For instance, Caggese and Cugnat (2008) make use of firm panel data to show that financially constrained firms use fixed-term workers more intensively than financially unconstrained firms do.

More in line with our paper, there is a strand of the literature focusing on the impact of the unions on the demand for temporary employment. The evidence offered by US studies is controversial. For instance, Abraham and Taylor (1996) argue that firms may use fixed-term contracts to counter the increase in labor costs brought about by unions representing the interests of permanent workers. However, Gramm and Schnell (2001), using a small sample of firms from Alabama, finds a negative correlation between union coverage and the probability of using temporary employees in the main occupational groups. Houseman (2001) uses a sample of more than 500 US firms and reaches the same conclusion. Conversely, using aggregate data from the US, Autor (2003) finds that the spread of temporary employment proved faster in states where the unions declined more slowly.

Clearly, evidence from the US, and from Anglo-Saxon countries more generally, does not necessarily constitute a suitable reference for our analysis, since the impact of unions and uncertainty on the firms' personnel policies depend fundamentally on the institutional architecture of the labor markets which, in turn, differs markedly between European and Anglo-Saxon countries.

As for Europe, the empirical research specifically investigating the effects of unions on firms' propensity to employ temporary workers is rather scant and provides mixed results. Francesconi and Garcia-Serrano (2004) find no evidence of a correlation between the share of temporary employment and unionization, using data on Spanish firms. Bryson (2007) uses data from British workplaces and reports a weak positive relationship between unions and the employment of agency workers. Salvatori (2012) also uses British data and finds no support for the hypothesis that firms under the threat of unionization are more likely to use fixed-term workers, and only weak evidence of negative effects for agency workers.

Of course, industrial relation practices and legislation on the utilization of fixed-term contracts differ markedly across the EU economies, which might partly explain the divergent results from

country to country. Salvatori (2009) provides the first comparative study by using workplace level data from a number of European countries. In general he reports that unionized workplaces are more likely to use temporary contracts across Europe, and hence the unions contribute to create labor market duality and welfare disparities between insider and outsider workers.

Aside from the difficulty of comparing country-specific evidence, another potential caveat of the empirical literature is that the impact of unionization on the firms' use of fixed-term contracts is generally examined without taking into account the role of economic uncertainty. This concern appears to be important for a number of reasons.

First, economic fluctuations have long been recognized as an important determinant of a firm's choice as to whether to hire on a temporary or permanent basis. For example, Bentolila and Saint Paul (1994) show that the difference in firing costs between permanent and temporary workers in Spain is associated with a higher cyclical elasticity (to sales) of the demand for temporary employment. In other words, firms can use temporary workers as a buffer, dropping them during recessions and allowing for a faster increase of employment during upturns. Boeri and Garibaldi (2007) find that, in the short run, Italian firms' propensity to hire on a temporary basis depends largely on the "states" of the economic conditions. Similar conclusions are reached by Holmlund and Storrie (2002), Costain et al. (2010), Nunziata and Staffolani (2007) and Lotti and Viviano (2011).

Second, including a measure of the firms' demand volatility is important to reduce omitted variable bias when investigating the impact of unionization on the demand for temporary employment. This is because output volatility is likely to be correlated with both the demand for temporary employment and union status if insider workers promote union activity to make their jobs safer.

4. Data and descriptive statistics

The empirical analysis is based on the *Employer and Employee Survey* (RIL) conducted by ISFOL in 2005 and 2007 on a nationally representative sample of firms operating in the non-agricultural private sector.

The RIL survey collects a rich set of information on personnel organization, industrial relations and other workplace characteristics. In particular, the RIL allows us to distinguish between the two different typologies of temporary contracts according as to whether a training clause is explicitly considered. We define as “training temporary contracts” (i.e., ‘*contratti a causa mista*’) any apprenticeship contracts, training and work contracts (*contratti di formazione lavoro*) and job insertion contracts (*contratti di inserimento lavorativo*). The term “non-training temporary contracts” (‘*Contratti a tempo determinato*’) refers to any other type of fixed-term contracts that contain no training clause.²

In order to link information on workplace characteristics to indicators of firm performance and economic volatility, the RIL dataset is merged with annual balance sheets and income statements data from the AIDA archive. The merge is carried out through unique company tax codes. As AIDA only covers limited companies with turnover over 100,000 euros, the RIL-AIDA merged sample restricts the analysis to limited companies above this (rather minimal) threshold.

Given our focus, we also exclude firms with fewer than ten employees, to retain only those firms characterized by a minimum level of organizational structure (and for which the share of employees in different type of contracts can be meaningfully computed). Our final sample includes over 6600 firms.

² The difference between these two types of temporary contracts appears crucial for firms’ personnel strategies. In fact, on-the-job training contracts are typically conceived as instruments to provide firm-specific training and an extended screening period to younger workers. Therefore, it is more likely that firms use these contracts for future hiring of workers on a permanent basis, rather than to exploit numerical flexibility. Indeed, for these contracts a “stepping stones” effect towards regular contracts is often found with empirical analyses of transition rates (Berton et al., 2011). Conversely, non-training fixed-term contracts are normally used by firms to cope with the numerical flexibility needed in a volatile economic environments, and generally offer workers poorer prospects of future transformation into permanent positions. These are in fact the type of contracts most often associated with “dead-end jobs” (Berton et al., 2011; Lilla and Staffolani, 2012).

Descriptive statistics on the pooled 2005-2007 RIL-AIDA sample are displayed in Table 1: the overall share of fixed-term contracts comes to 9.8%, of which the share of non-training contracts amounts to 6.1% and the share of training contract to 3.7%. As for the other variables of interest, we created a dummy variable (“union” for short) indicating whether workers have established any form of workers’ representation at the workplace that is legally entitled to participate in the firm-level bargaining process. Such workplace representation is present in about 40% of the firms. The table also shows our demand volatility measure, which has a mean equal to 0.4 in the sample. This volatility measure is computed as the average standard deviation of sales over the period 1997-2005 calculated at the 3-digit classification of economic activity (NACE, over 450 sectors). We used the AIDA dataset to compute this measure of volatility, as this dataset has yearly information on firms’ sales and number of employees (but not on unionization and contract types) for incorporated businesses in Italy. Note that our measure of demand volatility is computed at the sectoral level, rather than at the firm level, which helps minimize endogeneity concerns related to this variable.

Table 1 also sets out the descriptive statistics for all the other control variables used in the regression analysis. Table A1 in the Appendix provides definition of all the variables used in the analysis.

5. Econometric analysis

Our econometric analysis is performed with regression models to estimate the following equation:

$$(1) \quad TC_{it} = \beta_1 vol_{ct} + \beta_2 union_{it} + \beta_3 vol_{ct} \times union_{it} + X_{it} \delta + \varepsilon_{it} \quad \text{with } t=2005, 2007.$$

TC_{it} denotes the share of workers with a temporary contract in firm i at time t , vol_{ct} is our measure of a firm’s output demand volatility (in the sector c to which i belongs), and $union_{it}$ is a dummy variable indicating union presence at the workplace. $Vol_{ct} \times union_{it}$ is an interaction term capturing the way the presence of a union affects the impact of volatility on the share of fixed term contracts.

The vector X_{it} contains firm-level control variables that may affect the dependent variable, and may be correlated with union presence and volatility. In all the models reported below we have included the following control variables: the share of female workers, the shares of blue- and white-collar workers, the firm's age (in years), firm size dummies (3 classes defined in terms of employees), log value added per worker, sector dummies (10), geographical area dummies (4) and a dummy for year 2007. Experimentation with models containing sub-sets of these control variables never changed the estimates of our variables of interest in any appreciable way, and they are therefore not reported. Finally, ε_{it} is an idiosyncratic error whose properties are specified below.

5.1 Endogeneity and methodological issues

A major concern when estimating an equation like (1) concerns the potential endogeneity of a firm's union status. The presence of a union in a firm is influenced by many observed and unobserved firm characteristics, and by institutional features of the industrial relations operating at the national and local level. In particular, a number of variables might be expected to affect both a firm's union status and the propensity to hire temporary workers. Examples of such variables could be firm profitability or managerial culture and ability. Failing to control for them may potentially bias the main estimates of interest. For instance, a firm's culture that promotes high-performance resource management practices and high investment in human capital might generate both a limited use of temporary workers (particularly of the "buffer" type) and a cooperative environment conducive to worker representation, and hence union presence.

Assuming that omitted variables are time-invariant, consistent estimates may be obtained via linear models with fixed-effects. These models add individual-specific intercepts α_i as a component of the error term in equation (1), with α_i which can be freely correlated with union status or other RHS variables in the model. Since firms are observed in two points in time (2005

and 2007) in our data, we are technically able to estimate fixed effect models, albeit limited “within” variability and measurement errors are likely to make the conclusions arising from such estimates only suggestive.

Another solution to the potential endogeneity of union status is to resort to an instrumental variable strategy, which can potentially deal with time-invariant and time-variant omitted variable biases, including measurement error in union status. Of course a major challenge - still largely ubiquitous in the literature on the effects of union - is to find suitable instruments. Ideally, we would like to instrument union status in a first-differenced model. This empirical strategy would also cope with firm-level time-invariant unobserved heterogeneity (by eliminating the α_i) while controlling for time-specific omitted variables and measurement error in union status via IV. However, it is rather difficult to find valid instruments for the *change* in a firm's union status, and for this reason we had to abandon the panel IV approach.

Instead, we resort to a simpler IV framework, in which a firm's unions status is instrumented by (i) the two-year lagged mean industry unionization (at the three digit level), and (ii) the lagged industry profit per employee.³ Clearly, this IV strategy requires that we limit the sample to the 2007 cross-section, since our instruments are derived using the 2005 wave.⁴ We find that both instruments are strong predictors of current unionization.⁵ More importantly, we believe that both make for reasonable IV candidates. First, most temporary contracts are of very short duration, on average less than a year. Hence, a two-year lag in industry-level unionization is not expected to have any direct impact on the current use of temporary employment, once current unionization is

³ ‘Profit per employee’ is measured from the AIDA profit and loss accounts (after-tax accounting profits) and divided by the firm’s number of employees, and then averaged at the three-digit industry level.

⁴ No other dataset is currently available for Italy to compute mean unionization at a finely disaggregated level.

⁵ Salvadori (2009) reports IV estimates of the effect of unions on the probability that firms employ any temporary workers, using a 6-year lagged employment as an instrument for union status. The argument in favor of such an instrument is that lagged employment can be expected to be correlated with current union status, because union status is a persistent feature of the workplace and employment size is a known determinant of union status. However, our computations showed that long lags of employment (which are available from the information contained in AIDA) are only weak instruments in our application, with first-stage F statistics that never exceed critical values at conventional levels.

controlled for. While longer lags would, in principle, be desirable, we cannot construct them on the basis of the available data. Second, past firm profitability is a clear determinant of unionization, as unions can only expect to share rents where rents are sizeable. A two-year lag in profitability should suffice to remove concerns about possible direct effects on the current share of temporary workers.

Before turning to the results, we discuss the pros and cons of using linear models (as in our IV regressions) as opposed to nonlinear models such as Tobit regressions. The latter are a standard choice to account for the double censored nature of our dependent variable, the fraction of workers with a fixed-term contract lying between zero and one by construction (Houseman, 2001; Cappelli and Neumark, 2004; Lee and Kim, 2005). Linear models do not account for the censored nature of the dependent variable; on the other hand, however, they are better placed for dealing with the potential endogeneity of right-hand side variables and provide a simpler framework to compute average marginal effects in the presence of variable interactions (Angrist and Pischke, 2009).⁶ For these reasons, our main analysis is conducted through linear models. Nevertheless, for robustness we also report in the Appendix the results obtained with both standard and random effect Tobit models.

5.2 Econometric Results

Table 2 displays the estimated coefficients of equation (1), focusing, for the sake of brevity, only on the main variables of interest.⁷ The first column of Table 2 sets out the OLS coefficients of a simpler version of equation (1) that omits the interaction term. It shows that economic volatility has a positive and statistically significant effect on the utilization of flexible contractual arrangements, consistently with the previous literature arguing that volatility encourages firms to employ

⁶ Note that the existing IV Tobit models (see for instance IVTOBIT implemented in STATA) require that the endogenous variable be continuous, and hence are not suitable for dealing with the endogeneity of dummy-type RHS variables, as our union status indicator.

⁷ The estimated coefficients of the control variables are briefly discussed in the online appendix, where we compare our IV models with standard Tobit estimates.

temporary workers to adjust the labor input more easily. More specifically, an increase in volatility amounting to one standard deviation is associated with an increase of around 2 percentage points in the use of temporary contracts.

Column 1 also shows that the presence of a union in the workplace is associated with a lower use of temporary contracts (-2.3 percentage points). As argued above, one possible interpretation of this finding relates to the fact that temporary employees are less likely to be unionized and engaged in industrial relations than permanent employees. This implies that hiring workers on a temporary basis may significantly weaken the union bargaining power and its representativeness in industrial relations. According to this argument, the unions would discourage the use of fixed term contracts to favor outcomes for the “insiders” in future bargaining negotiations (Booth, 1995; Bentolila and Dolado, 1997).

To gain further insights into these mechanisms, we re-estimate equation (1) including the interaction term in column 2. In this case the unions no longer seem to play a direct role in discouraging the use of fixed term contracts: the estimated coefficient turns positive, though it is not statistically significant. However, it is the coefficient of the interaction term that is now negative and statistically significant. This implies that the effect of volatility is reduced by the presence of unions: a one standard deviation increase in volatility now raises the fraction of temporary employment in column 2 by about 1.88 percentage points (instead of 2 percentage points in the specification without interactions).⁸ Thus union opposition to the use of temporary employment seems to be driven by economic uncertainty: in a risky economic environment the unions may oppose an intensive use of fixed term contracts in order to prevent the weakening of their bargaining power. This is more likely to happen in volatile environments, characterized by high frequency of renegotiation of labor conditions and/or wage setting at the workplace.

⁸ This figure is obtained as $(0.122 - 0.07 * \overline{union}) * 0.2$, where $\overline{union}$ is the mean union in the sample, equal to 0.40 from Table 1.

Next, we consider a fixed effect (linear) model, as a robustness check that deals with omitted variable bias arising from time-invariant unobserved heterogeneity. Note that the results, set out in column 3, look quite similar to those shown by previous models. This is somewhat remarkable given the presence of only two waves and the consequent limited within-variability exhibited by both the fraction of temporary workers and unionization in the firm. This also suggests that the firms' unobserved heterogeneity does not seem to play a major role. Further, it is also interesting to note that controlling for time invariant unobserved heterogeneity has the effect of increasing the coefficient of the direct union effect, suggesting that firm unobserved heterogeneity, proxy for the quality of the firm, is positively correlated with union presence and negatively correlated with the share of temporary contracts, as expected.

The final part of Table 2 shows our IV estimates.⁹ Column (4) refers to the specification with no interaction between union and volatility. In this case, the availability of two instruments allow us to assess the validity of our IV strategy through a test of over-identifying restrictions. As shown by the Hansen J statistics, we are unable to reject the null hypothesis for the over-identifying restrictions, bearing out the validity of our instruments. The estimates in column (4) confirm the previous finding that volatility has a positive impact on the demand for temporary workers. Note that the direct union effect is now positive, though not statistically significant.

The specification in columns (5) include an endogenous interaction term, and the equation is exactly identified. The estimates largely confirm previous results: a positive impact of volatility and a negative interaction term between volatility and union presence. Additionally, the direct union effect is now positive and statistically significant.

⁹ In the IV models standard errors are clustered at the 3-digit industry level, as this is the level of aggregation of our instruments.

As a robustness check we include in the Appendix the estimates derived using standard and random effect Tobit models, which provide a consistent picture of the results obtained using linear models.

6. Contract type matters: training temporary contracts *vs.* non training temporary contracts

In this section we use our IV strategy to investigate the existence of any relevant heterogeneity across the two types of temporary contracts, and in doing so we are able to provide further insights for an understanding of the findings reported so far.

In Table 3, we show separate regressions for non-training temporary contracts (columns 1 and 2) and for training temporary contracts (columns 3 and 4). They clearly show that the effects of volatility and union presence are markedly heterogeneous across contracts types.

Considering the specification without the interaction term (columns (1) and (3)), the effect of volatility is large and statistically significant only for non-training temporary contracts, while it is close to zero and not statistically significant for training temporary contracts. This finding is consistent with the different role of the two contracts from the firm's standpoint. In coping with a volatile economic environment in the presence of a high level of protection for permanent workers, it is cost-effective for firms to rely on non-training temporary contracts rather than contracts with a training clause, as the training investment would be lost upon the temporary worker's dismissal.

The effect of unions is positive (and not significant) for non-training contracts and negative for training contracts. It can be rationalized on the basis of the different nature of the two types of contracts. Unions interested in protecting permanent insiders would encourage the hiring of those temporary workers that can work as a buffer, namely the non-training contracts. Instead, the training contracts cannot credibly work as a buffer to protect insiders, as firms will be less willing to sacrifice their training investments to accommodate the interests of the union.

Additional insights can be obtained on turning to models that interact unionization and volatility (column (2) and (4)). First, a positive and statistically significant direct union effect now emerges in the case of non-training temporary contracts. This effect might reflect both the interests of unions and firms. The unions have an interest in the presence of a positive share of the type of temporary workers that can offer a sort of protection to permanent workers. In turn, firms faced with an internal labor market dominated by the unions may want to hire on the cheaper type of temporary contracts to weaken the claims of regular (unionized) workers.

The sign of the interaction effect further qualifies such a finding, suggesting that, in a volatile economic environment, the unions become more concerned about the loss of bargaining power deriving from a high fraction of temporary workers and start opposing their presence. This additional effect is of appreciable magnitude: the estimated coefficients for the direct and the interaction terms imply that the effect of volatility is 0.24 for non-unionized firms, while only 0.03 for unionized firms. As already discussed, non-training temporary contracts are used by firms to cope with volatile economic environments. However, temporary workers are less prone to engage in industrial relations (typically organized by permanent workers), and their presence is thus expected to weaken the strength or “voice” of the unions in the bargaining process. This is conjectured to be the case mainly in highly volatile environments as employment conditions are negotiated more frequently in such environments. This effect can be quantified by computing the union effect at different values of volatility. Using the estimates in column 2, we then obtain that unions increase the proportion of non-training fixed-term workers by 7.5 percentage points when volatility is low, i.e. at the first decile of the volatility distribution. However, the same effect is rapidly reduced in the presence of higher volatility: it is only 0.06 if volatility is at the 25th percentile, and further drops to 0.04 at median volatility. Finally, the union effect is virtually zero

when volatility is very high (at the 90th percentile).¹⁰ It is also worth stressing that, even after accounting for the negative interaction term, the IV coefficients imply that the overall union effect is to increase the fraction of temporary employment in the firm, with the qualification that this fraction only refers to the non-training type.

As for the case of training temporary contracts, column 4 shows that neither the direct union effect nor the interaction effects are statistically significant. These findings reinforce our interpretation that training temporary contracts are used for purposes other than coping with economic volatility.

7. Conclusions

This paper has shown that workplace unionization has a positive effect on the firms' propensity to use temporary employment, although the effect is reduced by the presence of high economic volatility in the firm's output market. This conclusion, however, depends on the type of temporary contracts. Unions per se do generally encourage the use of non-training temporary workers, because they provide for a protecting buffer for permanent/insider workers. At the same time, firms may want to hire non-training temporary workers due to the presence of unions, as a way to reduce direct labor costs and labor turnover costs deriving from strict EPL for permanent workers. However, in the presence of high demand volatility, the unions become increasingly concerned about the loss of their bargaining power, and the union effect – while remaining positive – is reduced.

Conversely, workplace unionization does not exert any effect on the share of training temporary workers. We have argued that this finding can be accounted for by the different nature of this contractual arrangement. Training contracts are institutionally designed to favor more

¹⁰ Given the high significance of the underlying estimated coefficients, all these average treatment effects are also statistically significant.

efficient screening processes and investments in job-specific skills, rather than coping with the volatility in a firm's economic environment. Not surprisingly, employees on this type of contract have better chances of obtaining a permanent position in the medium run (Berton et al., 2011).

Overall, our paper provides some new insights into the role of worker representation at the workplace in affecting human resource management and firms' personnel policies. Moreover, our empirical evidence for Italy provides some additional support to the claims that the unions' preferences in the bargaining process tend to reflect the interest of the permanent workers, as traditionally formalized by insider-outsider literature.

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Table 1. Descriptive statistics

Variable	mean	std dev	min	max
share of temporary contracts	0.098	0.149	0	1
share of non-training fixed-term contracts	0.061	0.134	0	1
share of training temporary contracts	0.037	0.076	0	1
union	0.400	0.491	0	1
volatility	0.398	0.201	0.066	1.492
share of female	0.323	0.267	0	1
share of white collars	0.407	0.297	0	1
share of blue collar	0.539	0.313	0	1
share of trainees	0.200	0.316	0	1
firm age (years)	22.229	16.018	0	150
ln(value added per employee)	10.700	0.619	8.378	13.141
ln(n. of employees)	3.246	0.852	2.398	9.414
Macro-region				
North-West	0.368	0.482	0	1
North-East	0.260	0.439	0	1
Centre	0.196	0.397	0	1
South	0.177	0.381	0	1
Sector of activity				
quarrying, water, gas and electricity distribution	0.008	0.088	0	1
textile	0.129	0.335	0	1
manufacturing	0.180	0.384	0	1
mechanics	0.148	0.355	0	1
construction	0.161	0.368	0	1
trade, hotels and restaurants	0.172	0.377	0	1
transport and communication	0.040	0.196	0	1
Business services	0.073	0.260	0	1
education, health and public services	0.090	0.286	0	1
Year effects				
2007	0.447	0.497	0	1
N. of observations	6655			

Notes: sample weights used.

Table 2: OLS and IV models. All temporary workers.

	1	2	3	4	5
	Pooled OLS and FE models			2007 cross-section	
Variable	Pooled OLS	Pooled OLS	Fixed Effect	IV	IV
Volatility	0.101*** (0.019)	0.121*** (0.022)	0.122** (0.058)	0.157*** (0.029)	0.240*** (0.043)
Union	-0.023*** (0.004)	0.004 (0.009)	0.022 (0.016)	0.025 (0.051)	0.105** (0.053)
volatility*union		-0.069*** (0.024)	-0.071** (0.036)		-0.278*** (0.092)
Controls	yes	yes	Yes	yes	yes
N	6655	6655	5158	3328	3328
First-stage F stats:					
union				25.39	17.18
union* volatility					17.24
Exclusion restrictions:					
Hansen J statistics				0.175	0.010
Chi-sq(1) P-val				0.920	0.920

Notes: Coefficient estimates. IV models based on the 2007 cross section. In column 4 the instruments are the mean union status in 2005 at the 3-digit industry level and industry profits per worker in 2005 (at the three-digit). In column 5 the interaction term is also treated as endogenous and the additional instrument is the interaction between volatility and mean union status in 2005 at the 3-digit industry. Controls include: the share of female workers, the shares of blue and white collar workers, the firm's age, firm size dummies (3 classes defined in terms of no. of employees), sector dummies (10), area dummies (4). Standard errors in parentheses. Standard errors clustered at the 3-digit industry level in the IV models.

Table 3: Fixed-term contracts with and without training clause. IV models.

Variable	Non-training temporary contracts		Training temporary contracts	
	(1)	(2)	(3)	(4)
volatility	0.152*** (0.029)	0.205*** (0.041)	0.010 (0.011)	-0.011 (0.027)
union	0.066 (0.050)	0.117** (0.051)	-0.042* (0.022)	-0.068 (0.048)
volatility*union		-0.176** (0.088)		0.066 (0.068)
controls	yes	yes	yes	yes
N	3328	3328	3328	3328
First-stage F stats:				
union	25.4	17.18	25.4	17.18
union* volatility	-	17.24	-	17.24
Exclusion restrictions:				
Hansen J statistics	0.434	0.226	0.959	0.547
Chi-sq(1) P-val	0.510	0.634	0.327	0.460

Notes: Coefficient estimates. IV models based on the 2007 cross section. In columns 1 and 3 the instruments are the mean union status in 2005 at the 3-digit industry level and industry profits per worker in 2005 (at the three-digit). In column 2 and 4 the interaction term is also treated as endogenous and the additional instrument is the interaction between volatility and mean union status in 2005 at the 3-digit industry. Controls include: the share of female workers, the shares of blue and white collar workers, the firm's age, firm size dummies (3 classes defined in terms of no. of employees), sector dummies (10), area dummies (4). Standard errors in parentheses. Standard errors clustered at the 3-digit industry level.

Table A1: variable definition

log (value added)	Log of the valued added (source AIDA) and deflated by the value added deflator (source, ISTAT).
% Trained	Proportion of trained on total employment
log (n of employees)	Log of total number of employees
Union	Dummy variable that equals 1 if in the firm there is a worker representation on any kind, 0 otherwise
% Female	Proportion of female on total employment
% Managers	Proportion of managers and supervisors on total employment
% White-collars	Proportion of white collars on total employment
% Blue-collars	Proportion of manual as manual workers
% Fixed-term	Proportion of fixed-term workers on total employment
Volatility	average value of the standard deviation of sales over the period 1997-2005, calculated at the 3 digits classification of economic activity (NACE).
Macro-regions	4 dummies variables for: North Western regions, North-Eastern regions, Centre and Southern regions
Firms size	4 dummies variables for: n. employees<15 (ref. cat.), 14< n. employees<100, 99< n. employees<250, n. employees>249
Sectors	9 dummies variables for: Quarrying, gas, water and gas distribution (ref. cat.); textile; manufacturing; mechanics; Construction; retail and wholesale; transportation hotels and restaurants; insurance, monetary and financial intermediation; information, communication and other business services; health, education and other social services

Sources: AIDA and ISFOL-RIL Survey for 2005 and 2007

Appendix

A1. Robustness checks: Tobit regressions

In this appendix we include results obtained when using simple Tobit regressions to account for the double censored nature of our dependent variable. The error term is here assumed to be normally distributed and uncorrelated with the explanatory variables included in (1).

The results are shown in Table A2. Models 1 and 2 provide Tobit estimates for the pooled sample, i.e. waves 2005 and 2007, which we refer to as pooled Tobit. The upper part of the table shows the average treatment effects (ATE), namely the change in the share of fixed term contracts associated with a unit change in a given covariate, averaged over the sample distribution of the other covariates in the model. Note that, because of the nonlinear nature of the model, computing the ATE of a variable like union status requires that information be used on both the direct effect of union (coefficient β_2 in equation (1)) and the interaction effect (β_3 in (1)). The bottom part of the tables therefore also sets out the estimated coefficients for our main variables of interest.¹¹

Model 1 shows that economic volatility has a positive and statistically significant effect on the utilization of flexible contractual arrangements, with an estimated ATE at 0.049. This implies that a one standard deviation increase in volatility (equal to 0.2) raises the fraction of temporary workers in the firm by 1 percentage point ($=0.2 \times 0.049$). The estimates in Model 1 also show that the presence of a union in the workplace is associated with a lesser use of temporary contracts. Specifically, the ATE of union is a statistically significant reduction of 1 percentage point in the use of such contracts.

¹¹ It may be interesting to compare the estimated OLS coefficients derived in the paper with the related coefficients obtained with the Tobit models of Table A2. The first thing to observe is the close similarity in the two sets of estimated coefficients. However, it is worth noting that the magnitude of the estimated ATE is somewhat different. So, while volatility raises the fraction of temporary employment by 0.1 in the OLS models, the impact is reduced by roughly half in the Tobit model. On the contrary, the union impact is stronger in the OLS than in the Tobit model. These differences in part relate to the non-linear nature of the Tobit model; specifically, the Tobit ATE is computed conditionally on the dependent variable lying in the zero-one interval, whereas no such restriction is enforced in computing the OLS ATE. This caveat should be borne in mind also when observing the IV estimates.

We then estimate equation (1) with the inclusion of the interaction term (Model 2). The direct effect of unions is found to be positive, but not statistically significant. However, as in section 5, the coefficient of the interaction term is negative and significant.

In the rest of the Table, we also provide estimates of a random-effect Tobit model. In this case, an individual-specific random intercept α_i is added to equation (1) and is assumed to be normally distributed (with zero mean and variance σ_α) and uncorrelated with the other explanatory variables. Hence, this model produces consistent estimates only under the rather strong assumption that the α_i are uncorrelated with union status. This is the main reason that led us to prefer the linear FE model in the main text, as in that case the α_i are allowed to be freely correlated with union status. The magnitudes of the estimated coefficients in the random-effect Tobit are very similar to those discussed earlier for the pooled Tobit, suggesting that unobserved firm heterogeneity is a relatively minor concern in this application.

Table A2 also displays the estimates of other control variables that could affect the use of fixed-term contracts: the log of value added, firm size dummies, sectoral specialization, geographical location and workforce composition. It emerges that firm productivity (measured by the log of value added) is the only variable that is negatively correlated with the use of fixed-term contracts. On the other hand, firm size, the share of female workers, the share of low qualified workers (the omitted variable is constituted by executives) are positively associated with the use of total fixed-term contracts. The share of fixed-term contracts is also greater if firms are localized in North-East and Central regions. We do not include coefficients for industry dummies for the sake of space.

Table A2: Tobit regression estimates. All temporary contracts.

variable	Pooled Tobit						Random-effect Tobit					
	Model 1			Model 2			Model 3			Model 4		
	ATE	s.d.**	p-val.	ATE	s.d.**	p-val.	ATE	s.d.**	p-val.	ATE	s.d.**	p-val.
Volatility *	0.049	0.011	0.000	0.046	0.011	0.000	0.044	0.010	0.000	0.040	0.010	0.000
Union *	-0.010	0.003	0.000	-0.011	0.003	0.000	-0.009	0.003	0.001	-0.009	0.003	0.000
% female	0.031	0.006	0.000	0.031	0.006	0.000	0.034	0.005	0.000	0.034	0.005	0.000
% white collar	0.035	0.014	0.011	0.033	0.014	0.019	0.040	0.013	0.003	0.037	0.013	0.005
% blue collar	0.051	0.014	0.000	0.048	0.014	0.000	0.051	0.013	0.000	0.049	0.013	0.000
% trained	0.001	0.003	0.831	0.001	0.003	0.869	0.001	0.003	0.745	0.001	0.003	0.777
seniority	0.000	0.000	0.041	0.000	0.000	0.037	0.000	0.000	0.050	0.000	0.000	0.046
ln(value added)	-0.012	0.002	0.000	-0.012	0.002	0.000	-0.014	0.002	0.000	-0.014	0.002	0.000
14< n. employees<50	0.027	0.004	0.000	0.028	0.004	0.000	0.027	0.003	0.000	0.027	0.003	0.000
49< n. employees<250	0.050	0.005	0.000	0.049	0.005	0.000	0.048	0.005	0.000	0.047	0.005	0.000
n. employees>249	0.047	0.005	0.000	0.047	0.005	0.000	0.043	0.005	0.000	0.043	0.005	0.000
North East	0.009	0.003	0.001	0.009	0.003	0.001	0.009	0.003	0.002	0.009	0.003	0.002
Centre	0.018	0.003	0.000	0.018	0.003	0.000	0.019	0.003	0.000	0.019	0.003	0.000
South	0.002	0.004	0.659	0.002	0.004	0.624	0.002	0.004	0.657	0.002	0.004	0.621
Sector dummies	0.007	0.007	0.259	0.007	0.007	0.297	0.006	0.007	0.435	0.005	0.007	0.475
year 2007	-0.011	0.002	0.000	-0.011	0.002	0.000	-0.010	0.002	0.000	-0.009	0.002	0.000
sigma	0.183	0.004		0.183	0.004							
sigma_u							0.12	0.003	***	0.119	0.003	***
sigma_e							0.14	0.002	***	0.14	0.002	***
chi2							486.21			486.21		
Prob > chi2							0.000			0.000		
F(.)	13.39		12.89									
Prob > F	0.000		0.000									
Pseudo R2	0.433		0.442									
N of Obs	6655											
Estimated coefficients (selected variables)												
variable	Coef.	s.d.	p-val.	Coef.	s.d.	p-val.	Coef.	s.d.	p-val.	Coef.	s.d.	p-val.
volatility	0.112	0.025	0.000	0.143	0.029	0.000	0.099	0.022	0.000	0.127	0.024	0.000
union	-0.023	0.006	0.000	0.014	0.013	0.258	-0.020	0.006	0.001	0.016	0.013	0.222
union*volatility				-0.098	0.032	0.002				-0.092	0.030	0.002

Notes: standard errors clustered by firms.

* In models 2 and 4 the ATE for union and for volatility are computed after accounting for the interaction term in (1).

Controls include: the share of female workers, the shares of blue- and white-collar workers, the firm's age, firm size dummies (3 classes defined in terms of no. of employees), sector dummies (10), area dummies (4).



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