What Happened to the Italian Employment-Output Relationship?

Rebeca Jiménez-Rodríguez* — University of Salamanca
Giuseppe Russo† — University of Salerno

[Preliminary version, please do not quote]

June 6, 2007

Abstract

A well-established result of the literature on labour market rigidity is that the employment level is closely related to shocks in a deregulated labour market. This theoretical prediction has been rarely used to evaluate the effectiveness of labour market deregulation. This paper builds on a previous work where structural breaks in the employment series were found, and investigates how different employment indicators respond to aggregate shocks before and after the labour market deregulation started in the 90s. We find that the response to a shock of all our employment measures is substantially higher after the deregulation. This confirms that the Italian labour market has undergone important transformations in the last decades.

Keywords: labour market deregulation, impulse-response function.
JEL Classification: C22, J23

1 Introduction

One of the most important and established results of the literature on the effects of labour market regulation is that job protection does not entail *per se* a bad employment performance, its main effect being to smooth the path of the employment over the cycle (Nickell, 1978; Bentolila and Bertola, 1990; Bentolila and Saint-Paul, 1994). For job protection to affect the average employment, other distortions -like wage rigidities- are needed; moreover, its final effect is ambiguous (see, for example, Lazear 1990, or Bentolila and Bertola 1990). The

*Department of Economics. University of Salamanca. Campus Miguel de Unamuno. E-37007. Salamanca. Spain. Tel.: +34 923 29 46 40 (ext. 31 26). Fax: +34 923 29 46 86. E-mail: rebeca.jimenez@usal.es

†Correspondence to: Dipartimento di Scienze Economiche e Statistiche, Università di Salerno. I-84084. Fisciano. Italy. Tel.: +39 089 96 27 09. Fax.: +39 089 96 31 51. E-mail:grusso@unisa.it.

Italian labour market has been for decades one of the world's most regulated. In OECD (1994a, b) Italy was ranked first according to both employment protection and labour market regulation. After the rigidities set up in the 70s, we can observe the introduction of temporary apprenticeship contracts (1984) and the reduction of wage indexation (1986). Since 1991-92 we observe a wave of reforms which have deeply changed the labour market (OECD 2004).[1]

According to the theory, an effective deregulation is supposed to reinforce the link between employment volatility and macroeconomic volatility. That is, flexible and deregulated labour markets increase the responsiveness of the employment to macroeconomic shocks.

The aim of this paper is to investigate the Italian relationship between output and employment, relating it with the effectiveness of the Italian labour market deregulation. To do so, we first study the order of integration of all variables considering the existence of breaks around 1990-1993.[2] Once the order of integration of variables is established and in the case of non-stationarity of the levels of the variables, we test for the presence of a cointegration relationship between output and employment variables, considering also the breaks. Whereas a VEC model is estimated when evidence of cointegration is found , a VAR model is considered when there is evidence of no cointegration. We then calculate the impulse responses of employment variables to output shocks.[3]

The paper is organised as follows. Section 2 describes the methodology and our data. Section 3 reports the main findings. Section 4 presents the concluding remarks.

2 Methodology

To analyse the effects of output shocks on employment variables, we first study the order of integration of all variables considered in this study by performing unit root tests allowing for structural breaks. Specifically, we perform the test statistic S^{***} developed by Busetti and Taylor (2003). This test statistic accommodates for a possible break in both the intercept and the trend, and also allows for a variance shift. Once the order of integration of variables is established and in the case of non-stationarity of the levels of the variables, we test for the presence of a cointegration relationship between output and employment variables. To do so, we calculate the trace and maximum eigenvalue test statistics (see *e.g.* Johansen, 1995). The cointegration relationship between output and employment variables, however, may suffer regime shift and the standard cointegration tests tend to spuriously reject the null of no cointegration in this case. To solve such inconvenience, we also apply the Gregory-Hansen (1996) extension of the Engle-Granger (1987) test allowing for breaks in either the

[1] It is worth noting that this pattern has been similar in several European countries (OECD 2004).

[2] Jimenez-Rodriguez and Russo (2007) show the existence of breaks around 1990-1993 by using different measures of output and employment variables.

[3] Other aspects of the bivariate analysis, such as causality or asymmetry, are outside the scope of this paper.

intercept or the intercept and trend of the cointegrating relationship at an unknown time. In particular, Gregory and Hansen (1996) propose $ADF-, Z_{\alpha}-$, and $Z_t - type$ tests designed to test the null of no cointegration against the alternative of cointegration in the presence of a possible regime shift. In the case of evidence of cointegration, we use a VEC model. However, we consider a VAR model when there is evidence of no cointegration.[4] In both type of models (VEC and VAR) we the possible existence of breaks are taken into account.

We use quarterly data (1980:1-2003:1) of the following time series: Italian civilian employment (absolute value) (Source: ISTAT), Italian employment index (Source:OECD), real GDP (source: IFS), Standard units of labour (Source: ISTAT), Standard units of dependent labour, (Source: ISTAT).[5]

3 Empirical Results

We study the impact of an output shock on employment indicators. Prior to do so, we analyse the order of integration of the variables considered in this study using the test statistic S^{***} (Busetti and Taylor, 2003), which considers breaks in the intercept and the trend allowing for a change in the variance.[6] The results - shown in Table 1 - indicate the existence of a unit root for the level of the time series, but stationarity for the first log-differences. Given the evidence of non-stationarity found in the variables, we test for the presence of a cointegration relationship between output and employment variables. To do so, we apply the standard trace and maximum eigenvalue test statistics (see *e.g.* Johansen, 1995) and $ADF-, Z_{\alpha}-$, and $Z_t - type$ tests designed to test the null of no cointegration against the alternative of cointegration in the presence of a possible regime shift (see Gregory and Hansen, 1996). The results of tests are consistent with the lack of cointegration at the 5% significance level (see Tables 1 and 2). Therefore, we consider a bivariate VAR:

$$z_t = \beta_0 + \beta_1 z_{t-1} + ... + \beta_p z_{t-p} + \xi_t,$$

[4] For further discussion in this issue, see *e.g.* Hamilton (1994), and Ramaswamy and Slϕk (1998).

[5] The standard units of labour (SUL) are a measure of full-time equivalent labour input. They differ from other measures of employment because they are based on working hours, rather than on employment. For example, a worker may have two part-time jobs, thus he is worth one unit of full-time labour. Using the SUL gives us important additional information because they record not only if an individual is employed, but also if the adjustment to shocks is operated via the working hours. To compute these figures, ISTAT converts part-time jobs into equivalent full-time jobs according to some coefficient based on the average hours worked in a full-time job. The standard units of dependent labour (SUDL) differ from the SUL because they do not include self-employment.

[6] We consider the break dates reported in Jimenez-Rodriguez and Russo (2007) for the levels of the time series, and we estime the breaks for the first log-differences. The known breakpoints used to perform this statistic for the levels of the series have been: 1990:2 for overall output; 1993:1 for civilian employment; and 1992:3 for the rest of employment variables. The estimated breakpoints for the first log-differences are very nearby the known breakdates considered in the levels of the series.

where z_t represents a vector that contains an employment variable and an output variable. Given that all breaks are around 1990-1993, we split our sample into two subsamples: [1980:1-1992:3] and [1992:4-2003:1].[7] For each subsample, we estimate the model by Maximum Likelihood and assess the effects of output shock on employment variables.[8]

Figure 1 presents the generalised impulse response functions of employment indicators to output shocks before and after the break.[9] All employment variables react positively to a one unit shock to output, and the magnitude of response is larger in the post-1992 sample. This outcome gives further support to the Jimenez-Rodriguez and Russo (2007) findings.

4 Concluding remarks

In this paper we have tried to investigate the possible effect of labour market deregulation in Italy by using a time-series approach. Our idea is based on the well-established result that the transmission of aggregate shocks to the employment is stronger in labour markets with less job protection. In spite of its intuitiveness, this idea has not yet been used. Unlike other works that found only weak evidence for the effect of labour market reforms (see, for example, Auer 2005; L'Horty 2004; Winkelmann and Zimmermann 1998), our results find a quite important change in the dynamic behaviour of our employment measures after the deregulation. We argue that, indeed, these effects only appear when comparing the employment behaviour before and after the deregulation. From this institutional point of view, any attempt to assess the effect of deregulation should use a database going back to the 80s. Analyses based on the last decade are likely not to find any difference in the labour market behaviour, because the market regulation has been broadly constant over that period. Finally, our results seem to confirm the anecdotal evidence of increased job insecurity associated with short tenure and high worker turnover (for a measure of the increase in the gross worker turnover see Leombruni and Quaranta, 2005).

[7] We have chosen to split our sample in 1992:3 because it is the most frequent breakdate. However, our results are consistent with other dates around 1990-1993.

[8] The variables considered are I(1) in each subsample, but no evidence of cointegration is found.

[9] It is worth stressing that we analyse the responses of employment in industry to a one unit shock to industrial output.

References

[1] Auer, P. (2005). Protected Mobility for Employment and Decent Work: Labour Market Security in a Globalised World. *ILO Employment Strategies Papers* n.1.

[2] Bentolila, S. and Bertola, G. (1990), "Firing Costs and Labour Demand:How bad is Eurosclerosis?," Review of Economic Studies, 57, 381-402.

[3] Bentolila, S. and Saint-Paul, G. (1994), "A model of labor demand with linear adjustment costs," Labour Economics, 1, 303-326.

[4] Busetti, F. and Taylor, A. (2003), "Variance Shifts, Structural Breaks, and Stationarity Tests", Journal of Business and Economic Statistics, 21, 510-531.

[5] Engle, R.F. and Granger, C.W.J. (1987), "Cointegration and Error Correction: Representation, Estimation, and Testing", Econometrica, 55, 251-276.

[6] Gregory, A. W. and Hansen, B.E. (1996), "Residual-based tests for Cointegration in Models with Regime Shifts", Journal of Econometrics, 70, 99-126.

[7] Hamilton (1994), "Time Series Analysis", *Princeton University Press.*

[8] Johansen S. (1995), "Likelihood Based Inference in Cointegrated Vector Autoregressive Models". Oxford University Press, Oxford.

[9] Jiménez-Rodríguez, R. and Russo, G. (2007), "Institutional Rigidities and Employment Rigidity on the Italian Labour Market", mimeo, Universidad de Salamanca.

[10] L'Horty, Y. (2004). Instabilité de l'emploi : quelles ruptures de tendance?. *Document de Recherche EPEE* 04-01, University of Evry-Val d'Essonne.

[11] Lazear, E.P. (1990), "Job Security Provisions and Employment," Quarterly Journal of Economics, 105, 699-726.

[12] Leombruni R., Quaranta R. (2005). Eppure si muoveva già. Una breve storia della mobilità del lavoro in Italia. In Contini and Trivellato (Eds.), (2005) pp. 205-236.

[13] Nickell, S. (1978), "Fixed Costs, Employment and Labour Demand over the Cycle," Economica, 45, 329-345.

[14] OECD, (1994a). Employment Outlook, OECD PUBLICATIONS, 2, rue André-Pascal, 75775 PARIS CEDEX 16.

[15] OECD, (1994b), The OECD Jobs Study, OECD PUBLICATIONS, 2, rue André-Pascal, 75775 PARIS CEDEX 16.

[16] OECD, (2004), Employment Outlook, OECD PUBLICATIONS, 2, rue André-Pascal, 75775 PARIS CEDEX 16.

[17] Ramaswamy, R., and Slϕk, T. (1998), "The Real Effects of Monetary Policy in the European Union: What Are the Differences", *IMF Staff Papers* 2: 374-399.

[18] Winkelmann, R. and Zimmermann, K.F. (1998). Is Job Stability Declining in Germany? Evidence from Count Data Models. *Applied Economics*, 30, pp. 1413–20.

Table 1: *Busetti-Taylor stationarity tests*

			m_4	m_8
Levels	Break considered			
Civilian employment (abs. val.)	1993:2	S^{***}	0.365**	0.259*
Employment index	1992:3	S^{***}	0.450**	0.315*
Standard Units of Labour	1992:3	S^{***}	0.384**	0.285*
Standard Units of Dependent Labour	1992:3	S^{***}	0.538**	0.346**
Real GDP	1990:2	S^{***}	0.459**	0.316*
First differences	Break estimated ($\hat{\tau}_L$)			
Civilian employment (abs. val.)	1992:2	S^{***}	0.195	0.194
Employment index	1992:2	S^{***}	0.259*	0.254*
Standard Units of Labour	1992:2	S^{***}	0.254*	0.252*
Standard Units of Dependent Labour	1992:2	S^{***}	0.284*	0.253*
Real GDP	1990:1	S^{***}	0.229	0.280*

Note: The entries are the outcomes of the S^{***} stochastic stationarity statistic (developed by Busetti and Taylor, 2003) with known breakpoint for the variables in levels and with unknown breakpoint for the variables in first log-differences. In the latter case, parameter $\hat{\tau}_L$ denotes the breakdate estimate when a structural break occurs in the level and/or slope of the series - irrespective of whether or not there exists a break in the variance, and corresponds to the estimate of $S^{***}(\hat{\tau}_L)$. The null hypothesis of these tests is that the time series is stochastically stationary $[I(0)]$ versus the altenative hypothesis of the existence of a unit root $[I(1)]$, controlling for shifts in slope, level and variance. The statistics are computed using two values for lag-truncation parameter $m = \text{integer}(x(n/100)^{1/4})$, where x denotes the lags used to estimate the statistics and n the sample size used to calculate the statistic (for further details, see Busetti and Taylor, 2003). More concretely, we employ two values for x, namely, $x = 4$ and $x = 8$, which yield - for given n - m_4 and m_8, respectively.

One/two asterisks mean a p-value less than 5%/1%.

Table 2: *Standard cointegration tests*

$GDP-ICE$				5% *critical value*	1% *critical value*
	Trace statistic	*none*	5.337	15.41	20.04
		at most 1	2.316	3.76	6.65
	$Max-\mathrm{Ei}gen$ *statistic*	*none*	3.021	14.07	18.63
		at most 1	2.316	3.76	6.65
$GDP-IEI$					
	Trace statistic	*none*	6.092	15.41	20.04
		at most 1	2.131	3.76	6.65
	$Max-\mathrm{Ei}gen$ *statistic*	*none*	3.961	14.07	18.63
		at most 1	2.131	3.76	6.65
$GDP-SUL$					
	Trace statistic	*none*	4.264	15.41	20.04
		at most 1	0.914	3.76	6.65
	$Max-\mathrm{Ei}gen$ *statistic*	*none*	3.351	14.07	18.63
		at most 1	0.914	3.76	6.65
$GDP-SULD$					
	Trace statistic	*none*	9.138	15.41	20.04
		at most 1	2.556	3.76	6.65
	$Max-\mathrm{Ei}gen$ *statistic*	*none*	6.582	14.07	18.63
		at most 1	2.556	3.76	6.65

Note: For further details, see *e.g.* Johansen (1995).

Table 3: *Gregory and Hansen cointegration tests*

		Model C	*Model C/T*	*Model C/S*
$GDP-ICE$				
	ADF	-3.036	-2.999	-4.416*
	Z_t	-2.169	-2.884	-4.515*
	Z_α	-10.444	-16.995	-37.038*
$GDP-IEI$				
	ADF	-3.108	-2.669	-4.118
	Z_t	-2.098	-3.521	-4.240
	Z_α	-12.271	-23.320	-33.806
$GDP-SUL$				
	ADF	-2.820	-3.950	-4.383*
	Z_t	-2.711	-3.972	-4.464*
	Z_α	-17.606	-29.858	-35.505
$GDP-SULD$				
	ADF	-2.923	-2.588	-3.696
	Z_t	-1.877	-2.776	-3.947
	Z_α	-8.159	-17.795	-29.167

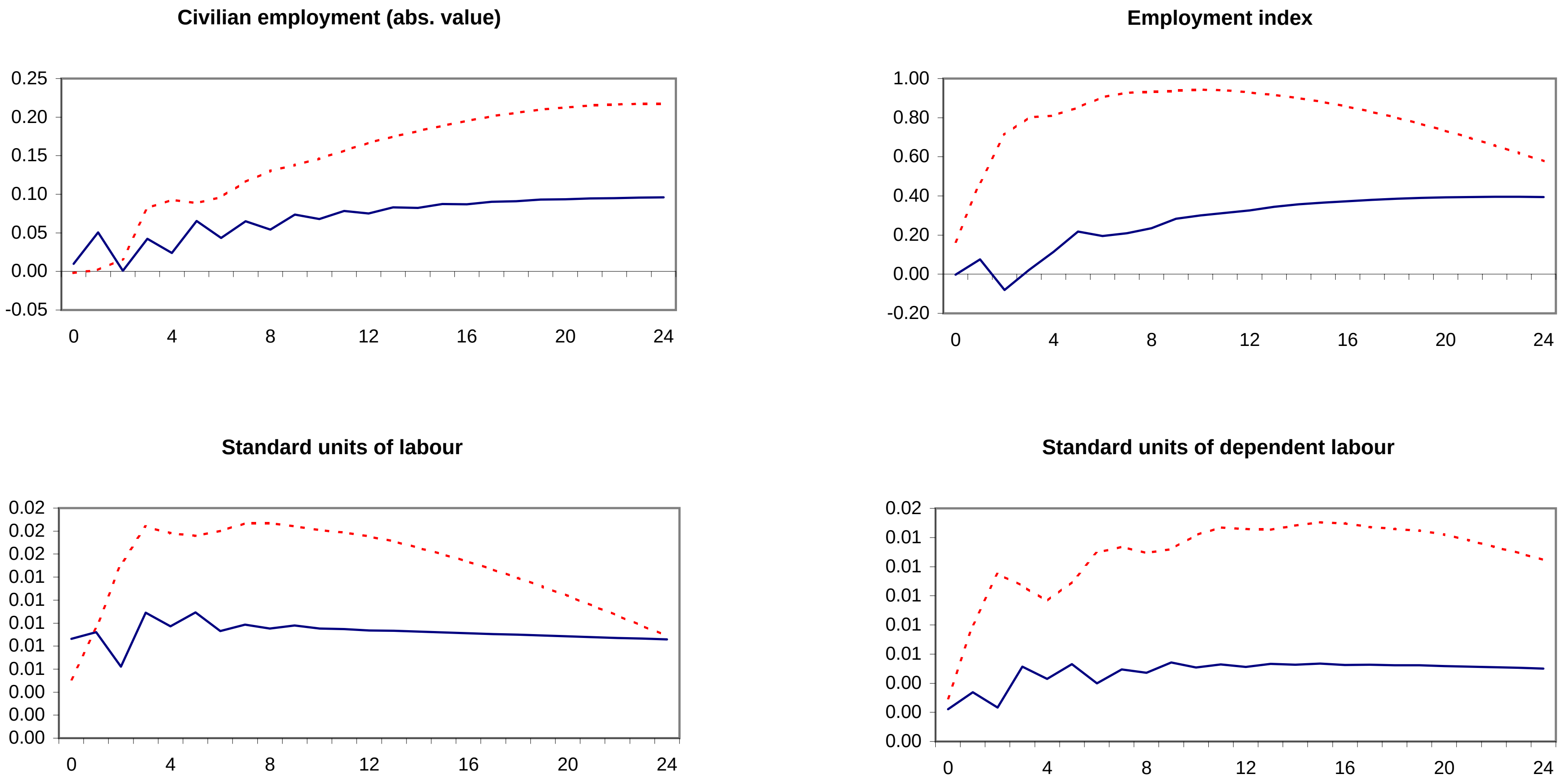
Note: The Gregory and Hansen (1996) cointegrating tests consider three models taking the structural change in the cointegrating relationship into account. Model C: A level shift in the cointegrating relationship can be modeled as a change in the intercept μ_0, with the slope coefficient α held constant (*level shift model*). Model C/T: A time trend can be introduced into the previous model (*level shift model with trend*). Model C/S: Another possible structural change allows the slope coefficient to shift as well (*regime shift model*). These models C, C/T, and C/S are estimated for each possible break date (for each τ), and the residuals $\hat{\varepsilon}_t$ are obtained. Next, a unit root test is performed on the estimated residuals, where the smallest values of the unit root test statistics are used to test the null of no cointegration against the alternative hypothesis of cointegration with a structural break. The authors consider the following Dickey-Fuller type-test (ADF^*) and Phillips type-test (Z_α^* and Z_t^*):

$$ADF^* = \inf_{\tau \in T} ADF(\tau),$$

$$Z_\alpha^* = \inf_{\tau \in T} Z_\alpha^*(\tau),$$

$$Z_t^* = \inf_{\tau \in T} Z_t^*(\tau).$$

Figure 1: Individual generalised impulse response function of employment indicators to a one unit shock to output



Note: Figure 1 presents the generalised impulse response functions of employment indicators to output shocks before (blue line) and after the break (red line)