Job Creation and the Self-employed Firm Size: evidence from Spain

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Abstract

Over the last decade, Spain created more jobs than any other country in Europe; however, now it is destroying them at an equal pace. Nevertheless, the relationship between the level of employment and the level of self-employment has remained relatively stable. In this context, there is a considerable policy interest in the way in which the self-employed firm in Spain creates and destroys employment, *i.e.*, in the role of the self-employed as creators of additional job opportunities. This paper provides evidence of the existence of a long-term relationship between the self-employed firm that hires external labor and employees hired by third parties in Spain, while accounting for the existence of an abrupt shift in the size of self-employed firms during the previous crisis (1991 to 1993). These findings are qualified testing whether this relationship is time-dependent. Our results suggest that the null hypothesis of linear cointegration would be rejected in favor of a two-regime threshold cointegration model, that is, in favor of a time-sensitive relationship with two opposite regimes. These two regimes differ in the way that the two components of the self-employed firm size respond to restore equilibrium. In this paper, alternative rationales for explaining these findings are also discussed.

Keywords: Self-employed firm size; Entrepreneurship; Job Creation; Cointegration; Self-employment; Structural Break; Threshold cointegration;

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1. Introduction

At the time of writing, the contribution of the self-employed sector is a topical policy issue given the extensive promotion of self-employment as a tool to combat unemployment, and as an alternative to the traditional active labour market policies.[1]
As is well-known, policies promoting self-employment may be considered promising in this respect.[2] In times of persistently high unemployment like the current one, self-employment is considered a means to combat unemployment given that new self-employed contributes to job creation not only directly, by creating jobs for the self-employed, but also indirectly, by the hiring of additional employees by self-employed firms (Pfeiffer and Reize, 2000a; Haltiwanger, 2006).[3] However, and leaving aside the problems related to converting unemployment into self-employment (Rissman, 2003; Baumgartner and Caliendo, 2008; Caliendo and Künn, 2010), more is not necessarily better (Baumol, 2008; Shane, 2009; Koellinger and Roessler, 2009; Congregado et al., 2010). For instance, the majority of self-employed firms are self-sufficient and consequently, only a minority hires other workers.[4] In addition, the coexistence of a map of incentives designed to foster self-employment and a labour market affected by stringent employment protection legislation (hereinafter EPL), in the context of a deep recession, may give rise to certain forms of "false" self-employment, promoting unexpected transitions from paid employment by third-party firms to self-employment (Román et al., 2009). The confluence of these three factors makes the Spanish economy a suitable case study. These factors could explain the minimal impact, on average, of the Spanish self-employed sector on job creation.

In addition, over the last decade, Spain created more jobs than any other country in Europe; however, now it is destroying them at an equal pace and has the highest unemployment rate in the European Union (20.2 % in late 2010). In this context, there is a considerable policy interest in the way in which the Spanish self-employed sector creates and destroys employment. The study of the relationship between the level of employment and the number of self-employed firms has therefore become crucial to

[1] In Calmforms, 1995, Pfeiffer and Reize, 2000a, 2000b and in Reize, 2004, among others, the role of self-employment promotion as an active labour market policy is discussed.

[2] These policies include loan guarantee schemes, technology-transfer and innovation programs, subsidized provision of business advice and assistance to small firms, and above all employment assistance programs (Parker, 2009).

[3] This indirect contribution to job creation is only done by self-employed who hires employees: employers or job creators. Both terms are used indistinctly in the text.

[4] In the EU-15 only 33,5% of self-employed are job creators (Eurostat, 2009).

understanding the current unemployment rate and the potential effectiveness of promoting self-employment as a way to combat unemployment.

However, research focusing specifically on the role of the self-employed as creators of additional employment opportunities is relatively scarce (see Carroll et al., 2000, Cowling and Taylor, 2001, as exceptions). In this context, the economics of self-employment should provide propositions and empirical evidence for understanding the reasons that the self-employed in certain countries or sectors have a greater propensity to employ additional workers. As Cowling (2003) states: "This is a significant gap in our knowledge and one which, when addressed, can provide policy-makers with valuable information to support the policy process."

This is precisely the goal of this article: to fill this gap, at least partially, by exploring the contribution of the self-employed sector to job creation in Spain, *i.e.*, the role of the self-employed as creators of additional employment opportunities, and analyzing the long-term relationship between self-employment and paid employment using a vector error correction model in which the error correction term is the self-employed firm size (hereinafter "*sfs*"), the core variable of our investigation.

The size, as usual, is measured by the number of employees divided by the number of self-employed firms. However, in order to capture the contribution of self-employment to job creation and to ensure that the self-employed firm size is not affected by the large group of solo self-employed firms (self-sufficient, "own account" workers), we redefine the self-employed firm size as the ratio of paid employees to job creators (*i.e.*, employers with employees).[5]

Although there is no consensus on the most appropriate measure of average firm size,[6] the aggregate average is usually obtained by dividing total employment by the total number of firms or self-employed people. However, if we consider solo self-employed firms in determining the net contribution of self-employment to paid employment, we may have an inaccurate measure for two reasons. First, it ignores certain aspects of solo self-employed firms (self-employed people who do not hire employees). These self-employed workers could be entrepreneurs of a "last resort" (Rissman, 2003) of a temporary nature, since they typically switch to paid employment when economic recovery occurs (Lucas, 1978). At this point, one could argue that their transitions from self-employment to paid-employment will alter the *sfs* with no real

[5] This term is defined by Cowling (2003).
[6] See Kumar et al. (1999) or Cabral and Mata (2003) for a detailed discussion.

effect on the total employment. Second, some of these own-account workers are undoubtedly true workers, dependent self-employed (Román et al. 2009), but official statistics do not account for the number of dependent self-employed. For these reasons, we need a measure of average firm size which is not affected by the large group of solo self-employed. In this paper, we propose the use of the paid-employment/employers ratio to circumvent these problems and to capture only the contribution of self-employment to paid employment.

In some aspects, our work is separate from but related to the large volume of work devoted to the analysis of the patterns of job creation at the level of the firm (Birch, 1979, Davis and Haltiwanger, 1992) or to the liability of newness (Short et al., 2009). Our work complements this type of dynamic study of job creation and destruction but is instead centered on the exploration of the long-term relationship between the self-employed utilizing additional labour and the number of jobs that they create.

In principle, a relatively low *sfs* might be the result of a high self-employment rate and/or of a weak ability of self-employed people to create employment. The overall size of the market (Smith, 1776, Becker and Murphy, 1992, Koellinger and Roessler, 2009), the regulatory framework (Garibaldi et al., 2004, Schivardi and Torrini, 2004), the availability of external funds (Laporta et al., 1997, Rajan and Zingales, 1998, Kumar et al., 1999), the sectoral composition (de Goey, 2004), the impact of tertiarization (Wennekers et al., 2010 or Bosma et al., 2009) and the use of more capital-intensive technologies (Lucas, 1978) could be considered explanatory factors of the *sfs* and its dynamics.

By using these arguments, one could argue that the entrance of Spain in the European Union (1986) should have given rise to radical changes in the Spanish self-employed sector: in particular, one should expect that new and thriving firms could have reaped the benefits of markets opening and embarked on creative or innovative ventures for commercial exploitation on a larger scale. However, the opening of the Spanish economy culminated in a resizing of both firms and sectors and in reallocations of labour with no significant changes in the average size of the self-employed firm.

Another argument that traditionally has been asserted in support of the negligible contribution of self-employment to employment in Spain is the adverse impact of EPL: in particular, the fact that more stringent labour laws tend to decrease the survival prospects of those entrepreneurs who employ outside workers (Parker, 2007). EPL imposes sunk costs for self-employed workers who decide to take on employees and, therefore, it may discourage them from employing additional workers if they think that their business will be prevented from reaching an optimal size (van Stel et al., 2007; Klapper et al., 2007). In these conditions, employers can circumvent this potential problem by contracting out in such a way that the self-employed firm size remains relatively stable, even in prosperous times.

This paper aims to investigate the interactions between paid employment, self-employment and self-employed firm size in the framework of a VECM model, using Spanish quarterly data during the period 1987:2-2008:4, analyzing empirically the relative stability of the Spanish *sfs*.

Our results show that, with the exception of one decline related to the crisis in the early 1990s, the average self-employed firm size has indeed been stable in Spain over the period from 1987 to 2008 in spite of the heavy fluctuations in job creation in Spain over the same period.

We also provide evidence of the existence of a long-term relationship between the self-employed who hire external labor and the wage-workers in Spain. These findings have qualified testing whether this relationship is time-dependent or not.

The structure of this paper is as follows. The second section shows some stylized facts and discusses previous evidence of the self-employed contribution to paid employment. The third section describes the data and the estimation methodology. The fourth section presents and discusses the results and performs a robustness check on the specification of the model. The final section concludes with a discussion of policy implications and some promising avenues for future research.

2. Self-employed firm size in the EU-15

Among other factors, self-employment firm size can be considered as an indicator of how entrepreneurs take advantage of the opportunities offered by the size of the market. Further, the extent of specialization is limited by the size of the market, *i.e.*, one could expect firm size to be correlated with the overall size of the market. In this sense, the forces of globalization or the European Single Market should have lead to an increase in the *sfs* in EU-15 economies. In the EU-15, there was a self-employed firm size of 5.9 in 2008 (Figure A1). The highest *sfs* were recorded in Luxembourg (14.9) and Denemark (10.78), followed by France, Sweden and Germany. At the other end of the range are Ireland, Portugal, and the three Mediterranean countries of Greece, Italy[7] and Spain. In particular, Greece recorded the lowest *sfs* (2.2). On the other hand, Austria and the countries that recorded an employment rate among employers above the EU-15 average also recorded a rate in excess of the EU-27 average (see figure A1).

Excluding own-account workers, the *sfs* expressed as the paid-employees/employers ratio (figure A3) reveals a quite similar picture. Employers in the EU-15 employed, on average, 17.7 employees in 2008. This figure varied considerably between Member States from highs of 30 persons per employer in Luxembourg and the UK and upwards of 23 in Sweden, Denmark, Netherlands and Finland to less than 15 in Ireland, Portugal,[8] and the Mediterranean countries.

With respect to the hire of additional workers, the Spanish self-employed sector compares relatively unfavorably to other countries. For instance, in 2008, the average number of paid-employees per employer was 14.4 (below the EU average) in spite of having created more jobs than any other country in Europe.

In the light of this trend we could argue that, twenty-five years after the entrance in the European Union (1986), the self-employed sectors in Spain appear to have failed to exploit the larger scales made available by the process of economic openness. The resizing of firms and sectors or the reallocations of labour have not lead to significant changes in the average self-employed firm size in spite of the great creation of jobs. To provide statistical evidence of this stylised fact, *i.e.*, the stability of the Spanish *sfs* in the long-run, will be the first objective of our empirical work.

Once the *sfs* stability is confirmed, we must analyze the response of the two components of the *sfs* to a shock, *i.e.*, how employees and self-employed react to restore the *sfs* equilibrium. This question is important given that its analysis provides useful information about the way in which employers create and destroy jobs.

Traditionally, the basic framework of Spanish economic relations, dating back to 1984, has been considered the main cause of the persistently high level of unemployment in Spain, where the unemployed are not properly incentivized to find employment and employers are not able to adapt to changing market conditions.[9] In

[7] In Santarelli and Vivarelli (2002) the role of public subsidies supporting the new firm foundation is analyzed in Italy.

[8] This stylized fact is present in the work of Baptista and Thurik (2007), who found a high proportion of "micro-businesses" created for subsistence which have little impact on growth and employment, in Portugal.

[9] The main feature of this framework was the low ability to reallocate labour between firms. Important institutional causes for this low ability include the high degree of employment protection corresponding

sum, from this perspective, employment protection legislation imposed sunk costs for self-employed workers who decided to hire employees by discouraging them from employing additional workers if they believed that their businesses would be prevented from reaching an optimal size. However, since 1984, the Spanish government has introduced five reforms in order to improve the performance of the Spanish labour market with a view to combating the rigidities described above. The first reform was introduced in 1984. Since the reduction of dismissal costs was politically damaging, the reform focused on facilitating the use of temporary employment contracts. As a result, a decade later, the percentage of temporary workers in Spain was the highest in Europe. The main concern with the liberalization of temporary contracts was that it generated segmentation between unstable low-paying jobs and stable high-paying jobs, without appearing to reduce unemployment. The resulting disparity between temporary and permanent workers was the impetus for two subsequent reforms (1994 and 1997), two true counter-reforms with the goal of reducing the high temporary rate.

The role of labour market institutions in the *sfs* can be also considered from the self-employed perspective. Although, as Robson (2003) and Torrini (2005) suggest, the relationship between self-employment and EPL may be negative[10], other authors suggest that this relationship may be positive.[11]

EPL imposes sunk costs for self-employed workers who decide to take on employees and, therefore, it may discourage them from employing additional workers if they think that their businesses will be prevented from reaching an optimal size (van Stel et al., 2007; Klapper et al., 2007). In these conditions, employers can circumvent this potential problem by contracting out in such a way that self-employed firm size remains relatively stable, even in booming times, *i.e.*, increases in paid employment are also associated with increases in self-employment-.

Previous discussion allow us to establish three hypotheses:

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to a low level of labour market flexibility; the importance of collective bargaining to establish employment conditions; the low level of functional and geographical mobility reinforced by the need to acquire court's approval for changing job's functional and geographic characteristics (García-Serrano and Jimeno, 1998; Dolado and Jimeno, 1997); and the generosity of the Spanish unemployment benefit system discouraging the search for employment (Bover et al. 2002, Blanchard and Jimeno, 1995)

[10] Because a higher EPL implies a higher opportunity cost of being entrepreneur.

[11] For instance, a combination of a strict EPL together with schemes oriented to encourage people to start business might give rise to the establishment of mutual agreements between employers and employees, just to evade the most onerous elements of employment legislation giving rise to the so-called phenomenon of 'dependent' self-employment. Thus, these employees are 'pushed' to self-employment, although doing the same activity, taking advantage of incentives schemes and reducing tax liabilities. As a result, a positive relationship between self-employment and EPL is suggested. (Román et al. 2009).

i) There is a long-term relationship between the two components of *sfs*.

ii) Stability: If Spanish employers circumvent the employment protection legislation by contracting out, increases in paid-employment do not lead increases in the *sfs*.

iii) Volatility: Spanish self-employed workers who decide to take on employees as rational agents will try to substitute permanent workers with temporary workers. As a result, more volatility in paid-employment should be observed.

Thus, an analysis of the relationship between employees and job creators should be performed in order to ascertain the effect on the *sfs* of the evolution of its two components and by the labour market institutions.

3. Econometric methodology and data

As we mentioned, the core variable of our investigation is the self-employed firm size. To carry out this task, we analyze the relationship in the long run between the two components of the *sfs*, wage workers and employers. In so doing, we look for robustness using alternative and complementary econometric methods.

As a preliminary step, we examine the time series properties of the two variables before analysing the linear long-run cointegration between them. However, it is well-known that one of the most likely reasons for the rejection of a linear long-term relationship between two series is the presence of nonlinearity. We then account for nonlinearity by applying a threshold cointegration. The concept of threshold cointegration characterizes a discrete adjustment, in a way in which the cointegration relationship between a set of variables exists only in a certain range but does not hold if the system deviates too much from the equilibrium. Hansen and Seo (2002) provide a vector error-correction model (VECM) in which exist a cointegration relationship between both variables and a threshold effect as an error correction term. [12] One of the most interesting points of our estimates emerges with respect to the interpretation of the error correction term, since it can be interpreted as the log of the *sfs*. After employing a non-linear econometric methodology, we verify the robustness of our model by using the Bai-Perron (1998, 2003) approach to detect the presence of structural breaks in the

[12] These authors also develop a test LM in order to test the threshold effect presences in the cointegration relationship.

sfs, *i.e.*, testing whether the break date and the date corresponding to the threshold are identical.

The data used in the empirical analysis are quarterly observations drawn from the Labour Force Survey (EPA) conducted by the Spanish National Institute of Statistics (INE). The sample period is from 1987:2 to 2008:4. Our variables are defined to exclude the agricultural sector.[13] As discussed in the introduction above, the *sfs* is defined as the ratio of private wage earners to employers, using the term employer or job creators to define the subset of the self-employed who hire workers –employers with employees or job creators.[14]

4. Results

4.1. Tests for unit roots

As a preliminary step in our analysis, we examine the time series properties of private wage earners and employers by testing for a unit root over the full sample. We have used a modified version of the Dickey-Fuller and Phillips-Perron tests proposed by Ng and Perron (2001), which try to solve the main problems present in these conventional tests for unit roots.[15] This method consists of a class of modified tests, called $\overline{\boldsymbol{M}}^{GLS}_{MAIC}$, originally developed in Stock (1999) as *M* tests with GLS detrending of the data as proposed in Elliot et al. (1996) and using the Modified Akaike Information Criteria (*MAIC*).[16] In addition, Ng and Perron (2001) have proposed a similar procedure to correct for the deficiencies of the standard Augmented Dickey-Fuller (ADF) test, $\boldsymbol{ADF}^{GLS}_{MAIC}$.[17]

The results of Ng and Perron tests are illustrated in Table 1. As shown in the table, the existence of two unit roots is clearly rejected at the usual significance levels for all variables, whereas the null hypothesis of nonstationarity in levels is clearly rejected at the usual significance levels for the employers variable but cannot be

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[13] Workers in the agricultural sector are excluded because this sector is structurally different from the rest of the economy. The availability of non-agricultural employment data, constrains the starting point of our data, since it is not possible to disaggregate agricultural and non-agricultural workers using Spanish statistics before 1987:1.

[14] The relation between the two variables in logs is given by $\boldsymbol{w_t} - \boldsymbol{e_t} = \ln(\boldsymbol{W_t} / \boldsymbol{E_t})_t$ where $(W_t / E_t)_t$ is the *sfs*.

[15] See, Dejong et al., 1992; Schwert, 1989 and Perron and Ng, 1996.

[16] These tests are the $\overline{\boldsymbol{MZ}}^{GLS}_{\alpha}$, $\overline{\boldsymbol{MSB}}^{GLS}_{\alpha}$ and $\overline{\boldsymbol{MZ}}^{GLS}_{t}$.

[17] See Ng and Perron (2001) and Perron and Ng (1996) for a detailed description of these tests.

rejected for the private wage earners variable (at the 5% significance level). Thus, according to the results of these tests, w_t would be I(1), but e_t could be I(1) or I(0).

Table 1
Ng and Perron[a,b] tests for a unit root

I(2) vs. I(1)	Case: p = 0, $\bar{c} = -7.0$			
Variable	$\overline{M}^{GLS}_{MAIC}$ tests			
	$\overline{MZ}^{GLS}_{\alpha}$	$\overline{MZ}^{GLS}_{t}$	$\overline{MSB}^{GLS}_{\alpha}$	ADF^{GLS}_{MAIC}
Δw_t	-26.55***	-3.31***	0.124	-4.15***
Δe_t	-41.95***	-4.57***	0.109	-10.30***

I(1) vs. I(0)	Case: p = 1, $\bar{c} = -13.5$			
Variable	$\overline{M}^{GLS}_{MAIC}$ tests			
	$\overline{MZ}^{GLS}_{\alpha}$	$\overline{MZ}^{GLS}_{t}$	$\overline{MSB}^{GLS}_{\alpha}$	ADF^{GLS}_{MAIC}
w_t	-1.53	-0.87	0.565	-0.89
e_t	-18.33**	-2.99**	0.163**	-3.41**

Notes:
[a] A *, ** and *** denote significance at the 10%, 5% and 1% levels, respectively;
[b] The *MAIC* information criteria is used to select the autoregressive truncation lag, k, as proposes in Perron and Ng (1996). The critical values are taken from Ng and Perron (2001), table 1.

Critical values:	Case: p = 0, $\bar{c} = -7.0$			Case: p = 1, $\bar{c} = -13.5$		
Variable	10%	5%	1%	10%	5%	1%
$\overline{MZ}^{GLS}_{\alpha}$	-5.7	-8.1	-13.8	-14.2	-17.3	-23.8
$\overline{MSB}^{GLS}_{\alpha}$	0.275	0.233	0.174	0.185	0.168	0.143
, $\overline{MZ}^{GLS}_{t}$ ADF^{GLS}_{MAIC}	-1.62	-1.98	-2.58	-2.62	-2.91	-3.42

A potential difficulty in assessing the time series properties of employment variables is that they can be subject to potential structural breaks. Hence, in order to provide further evidence on the degree of integration of e_t, we have also applied the Perron-Rodriguez test for a unit root in the presence of a one-time change in the trend function. Perron and Rodriguez (2003) extend the tests for a unit root analyzed by Perron and Ng (2001) to the case where a change in the trend function is allowed to occur at an unknown time, T_B. We use the method where the break date is selected to maximize the absolute value of the t-statistic on the coefficient of the change in slope as suggested by Perron (1997). The results are presented in Table 2. We consider the Model II where a structural change in intercept and slope is allowed to occur at an unknown time. Using the *MAIC* to select k, there is no evidence against the unit root for the employer series. Thus, according to the results of these tests, e_t would be I(1).

Table 2
Perron and Rodríguez[a,b] tests for a unit root with one time change in the trend function choosing the break point maximizing the $\left|t_{\hat{\beta}_2}\right|$

I(1) vs. I(0)	Case: p = 1, $\bar{c} = -22.5$					
Variable	$\overline{M}_{MAIC}^{GLS}$ tests					
	$\overline{MZ}_{\alpha}^{GLS}$	$\overline{MZ}_{t}^{GLS}$	ADF_{MAIC}^{GLS}	k	T_B	$\hat{\alpha}$
e_t	-9.56	-2.14	-2.40	6	1992:4	0.74

Notes:
[a] A *, ** and *** denote significance at the 10%, 5% and 1% levels, respectively.
[b] The MAIC information criteria is used to select the autoregressive truncation lag, *k*, as proposed Perron and Rodriguez (2003). We impose a minimal value *k* = 1. The critical values are taken from Perron and Rodriguez (2003), table 1 (a), Model II, T = 100.

Critical values:	Case: p = 1, $\bar{c} = -22.5$		
	10%	5%	1%
$\overline{MZ}_{\alpha}^{GLS}$	-20.7	-22.9	-27.0
$\overline{MSB}_{\alpha}^{GLS}$	-3.19	-3.35	-3.66
, ADF_{MAIC}^{GLS}	-3.59	-3.83	-4.31

4.2. Exploring the long-run relationship between employers and employees

We estimate the long-run regression model using the Dynamic Ordinary Least Squares (DOLS)[18] estimation method of Stock and Watson (1993), extended by Shin (1994).[19] Shin (1994) approach is similar to the KPSS[20] tests, which, for the case of cointegration, are implemented in two stages.

The first step in our estimation strategy would therefore consist of the estimation of a long-run dynamic equation including leads and lags of the explanatory variables in the long-run regression model, *i.e.*, the so-called DOLS regression:

$$w_t = \delta + \beta e_t + \sum_{j=-q}^{q} \varphi_j \Delta e_{t-j} + \varepsilon_j \quad (1)$$

[18] LS estimation of equation might suffer two problems: nuisance parameter dependences due to serial correlation in the residuals and endogeneity bias arising from innovations in employees to innovations in employers.
[19] In order to overcome the problem of the low power of classical tests for cointegration under the presence of persistent roots in the residuals of the cointegration regression, Shin (1994) suggested a new test where the null hypothesis is cointegration.
[20] These tests are called the Kwiatkowski et al. (1992) tests, and assume the null hypothesis of stationarity.

The second step is to use the statistic C_{μ}, a LM-type test designed by Shin (1994), to test the null of cointegration against the alternative of no cointegration in a DOLS regression.[21] In Table 3, we report the estimates from the DOLS regression and the results from Shin's test. We see evidence of linear cointegration between w_t and e_t, because we do not reject the null hypothesis of cointegration, being the estimated value of the paid-employment elasticity of employers, β= 0.83 with an a priori expected positive sign.

Table 3
Stock –Watson-Shin's DOLS [a,b,c,d] estimation of linear cointegration

Parameter estimates	1987:2-2008:4 Full sample
$\hat{\delta}$	3.55 (3.21)
$\hat{\beta}$	0.83 (5.10)
Test:	
$C^{c}{}_{\mu}$	0.200
R^2	0.99
$\hat{\sigma}^2$	0.086

Notes:
[a] *t*-statistics in brackets. Standard Errors are adjusted for long-run variance. The long-run variance of the cointegrating regression residual is estimated using the Barlett window which is approximately equal to $INT\left(T^{1/2}\right)$ as proposed in Newey and West (1987).
[b]We choose $q = INT\left(T^{1/3}\right)$ as proposed Stock and Watson (1993).
[c] C_{μ},is a LM statistic for cointegration using the DOLS residuals from deterministic cointegration, as proposed Shin (1994). A *, ** and *** denote significance at the 10%, 5% and 1% levels, respectively.
[d]The critical values are taken from Shin (1994), table 1, from m=1:

Critical values:	10%	5%	1%
C_{μ},	0.231	0.314	0.533

4.3. Modelling non-linearity

Results obtained in the previous section point to the existence of a long-run relationship between employers and wage workers. However, there is no reason to assume a symmetrical relationship, as a prior. As mentioned above, it is possible to have a time-varying relationship. We should therefore take into account the possible existence of asymmetries in the relationship. Our benchmark linear model is a finite-order VAR of the following form:

[21] C_{μ} is the test statistic for deterministic cointegration, i.e., when no trend is present in the regression.

$$x_t = c + \sum_{i=1}^{k} A_i x_{t-i} + \varepsilon_t \quad (2)$$

In the above model, $x_t = [w_t, e_t]'$ is a vector of non-stationary variables containing the log of wage earners (w_t) and the log of employers (e_t), A_i is a *2x2* matrix of parameters, and ε_t is a *2x1* vector of residuals. In order to characterize the long-run dynamic adjustments, we can rewrite the equilibrium VAR model as a vector error correction model (VECM):

$$\Delta x_t = c + \sum_{i=1}^{k-1} \Gamma_i \Delta x_{t-i} + \Pi x_{t-k} + \varepsilon_t \ (3)$$

where $\Gamma_i = -\left(I \ - \sum_{i=1}^{k-1} A_i \right)$ and $\Pi = -\left(I - \sum_{i=1}^{k} A_i \right)$. The matrix Π is usually decomposed as $\Pi = \alpha\beta'$, where α and β are *n* x *r* matrices containing the adjustment coefficients and the cointegrating vector, respectively, *n* is the number of variables, and *r* is the number of cointegrating relationships. The symbol Δ in equation (3) is the first difference operator. In this form, all terms in equation (3) are stationary, that is, integrated of order zero. In our application the system can be written as:

$$\begin{bmatrix} \Delta w_t \\ \Delta e_t \end{bmatrix} = \Gamma(L) \begin{bmatrix} \Delta w_{t-i} \\ \Delta e_{t-i} \end{bmatrix} + \begin{bmatrix} \alpha_w \\ \alpha_e \end{bmatrix} (w_{t-1} - \beta e_{t-1}) + \begin{bmatrix} \varepsilon_t^w \\ \varepsilon_t^e \end{bmatrix} (4)$$

where α_w and α_e indicate the speed of adjustment of each variable back to its long-run value. In the above model, the lagged residuals from the cointegrating vector act as an error correction term. This term, $w_{t-1} - \beta e_{t-1}$, captures the extent of disequilibrium for the system of variables with respect to the long-run relation between all variables in the system. A significant error correction term (*i.e.*, a significant α parameter) implies long-run causality from the explanatory variables to the dependent variable under consideration. However, there is no reason for assuming a symmetrical relationship between wage earners and employers, as a prior. Our model should therefore allow for the possibility of a nonlinear relationship. We account for non-linearity by applying a threshold cointegration model (Balke and Fomby, 1997) as a feasible means of combining non-linearity and cointegration. We can use this approach to allow for a non-linear adjustment to the long-run equilibrium. Hansen and Seo (2002) provide a vector error-correction model (VECM) in which a cointegration relationship exists between two variables and a threshold effect as an error correction term. As an extension of model (4), a two-regime threshold cointegration model takes the form

$$\begin{bmatrix} \Delta w_t \\ \Delta e_t \end{bmatrix} = \Gamma(L) \begin{bmatrix} \Delta w_{t-1} \\ \Delta e_{t-1} \end{bmatrix} + \begin{bmatrix} \alpha_{11} \\ \alpha_{21} \end{bmatrix} (w_{t-1} - \beta e_{t-1}) + \begin{bmatrix} u_t^w \\ u_t^e \end{bmatrix} with\ (w_{t-1} - \beta e_{t-1}) \le \gamma$$

$$\begin{bmatrix} \Delta w_t \\ \Delta e_t \end{bmatrix} = \Gamma'(L) \begin{bmatrix} \Delta w_{t-1} \\ \Delta e_{t-1} \end{bmatrix} + \begin{bmatrix} \alpha'_{11} \\ \alpha'_{21} \end{bmatrix} (w_{t-1} - \beta e_{t-1}) + \begin{bmatrix} v_t^w \\ v_t^e \end{bmatrix} with\ (w_{t-1} - \beta e_{t-1}) > \gamma \quad (5)$$

where γ is the threshold parameter that delineates the two different regimes. Using model (5) as a basis, the first step of our analysis consists of testing whether the dynamic behaviour and the adjustment towards the long-run equilibrium are linear or exhibit a threshold. Hansen and Seo (2002) propose a set of heteroskedastic-consistent Lagrange Multiplier (LM) test statistics for the null hypothesis of linear cointegration (*i.e.*, there is no threshold effect) against the alternative of threshold cointegration (*i.e.*, model 5). The results of the test are reported in Table 4. Threshold cointegration would appear at the 3% significance level such that the null hypothesis of linear cointegration is strongly rejected.

Table 4
Hansen-Seo tests of threshold cointegration

	$\sup \mathrm{LM}^0\ \beta = 1$	$\sup LM\ \hat{\beta}$ *estimated*
Test statistic value	23.708	21.937
Bootstrap p-values	0.030	0.072
Fixed regressor p-values	0.044	0.004
Threshold parameter	2.657	2.860
Estimate of the cointegrating vector	1	0.969

Once the presence of threshold effects is confirmed, we then estimate model. The estimated threshold is $\hat{\gamma} = 2.657$, when β is fixed.[22] Hence, the first regime (including 82.14% of the observations) would occur when the self-employed firm size is below 14.25 employees per employer; in other words, when the gap (*sfs*, in logs) is below 2.657. In turn, the second regime would occur when the *sfs* is above 2.657. This is a relatively unusual regime, including 17.86% of the observations (1987:2 to 1991:3). The corresponding two-regime threshold VECM results are presented in Table 5, where

[22] We consider the results for the fixed beta slightly more accurate for interpretation and, we focus on these results from now on. In that sense β does not significantly differ from 1, when β is estimated, which allows us to fix the value of β at 1. In this way the error-correction term equals $w_{t-1} - e_{t-1}$, i.e. the error correction term is equal to the logarithm of the self-employed firm size. In any case, estimates with non-fixed β are reported in the appendix B –table B1-.

significant error-correction effects appear in the second regime but not in the first regime for the employers´ equation. For the private wage earners´ equation the adjustment coefficient is significantly negative when the *sfs* is below 2.657 (regime 1), implying that a value of the *sfs* below 14.25 in one quarter produces downward pressure on private wage earners to restore the long-run equilibrium in the next quarter. By contrast, in this regime, no such pressure exists, as the error-correction term in the employers' equation is not significant. As α_{21} is not statistically different from zero, employers is said to be long-run weakly exogenous with respect to the long-run equilibrium in this first regime, *i.e.*, a lack of dependence of e_t on *sfs* shocks. As regards the second regime, the adjustment coefficients α'_{11} and α'_{21} are significantly different from zero when the *sfs* is above the threshold and the effects are both positive.

Table 5
Threshold VECM Estimates, fixed beta (Hansen-Seo, 2002)

Threshold	Regime 1 $(w_{t-1}-e_{t-1}) \leq 2.657$ $sfs \leq 14.25$		Regime 2 $(w_{t-1}-e_{t-1}) > 2.657$ $sfs > 14.25$	
Variable	Δw_t	Δe_t	Δw_t	Δe_t
C	0.278	-0.044	-0.516***	-0.603*
	(0.084)	(0.151)	(0.126)	(0.339)
Δw_{t-1}	0.407***	-0.080	-0.483***	-0.020
	(0.137)	(0.323)	(0.172)	(0.253)
Δw_{t-2}	-0.205***	0.039	-0.772***	-1.058***
	(0.081)	(0.220)	(0.155)	(0.432)
Δe_{t-1}	-0.041	-0.155	0.193***	-0.174
	(0.067)	(0.105)	(0.077)	(0.358)
Δe_{t-2}	-0.083	-0.096	0.016	0.663*
	(0.068)	(0.145)	(0.070)	(0.190)
α	-0.105***	0.024	0.195***	0.226**
	(0.033)	(0.059)	(0.046)	(0.122)
Observations percentage:	82.14 % [1991:3-2008:4]		17.86 % [1987:2-1991:3]	

Notes:
Standard errors in brackets
*, **,*** Significant at the 10%, 5% and 1% levels, respectively.

As is well-known, for values of α close to unity, the disequilibrium will be entirely eliminated within one quarter. For 0<α<1, the dynamic adjustment path will be monotonically convergent.

The absolute value of α, gives information about the number of quarters needed to restore the long term equilibrium.[23] Table 6 shows that, in the first regime, the current one, after almost $1^{1/2}$ years 50% of the disequilibrium gap created by the shock has been closed by the adjustment in paid employment. The speed of adjustment is now

[23] In order to obtain the number of quarters (t), required to dissipate x% of a shock we must compute $(1-\alpha)^t=1-x\%$, where α is the absolute value of the estimated speed adjustment parameter.

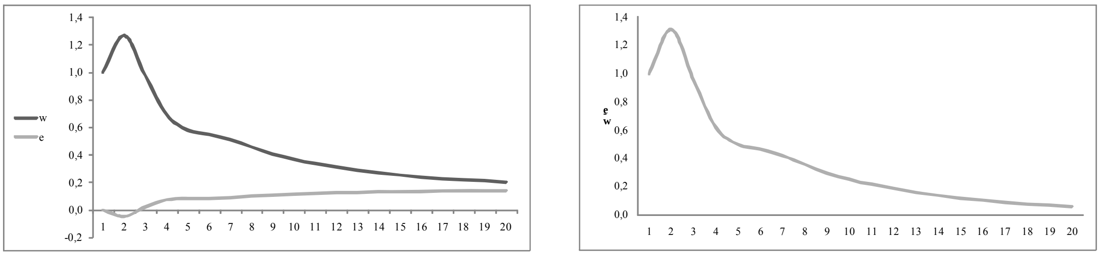
lower than before: in the past –regime 2- it takes less than one year to close the gap created by an equilibrium distorting shock.

Table 6
Estimated Speeds of Adjustment

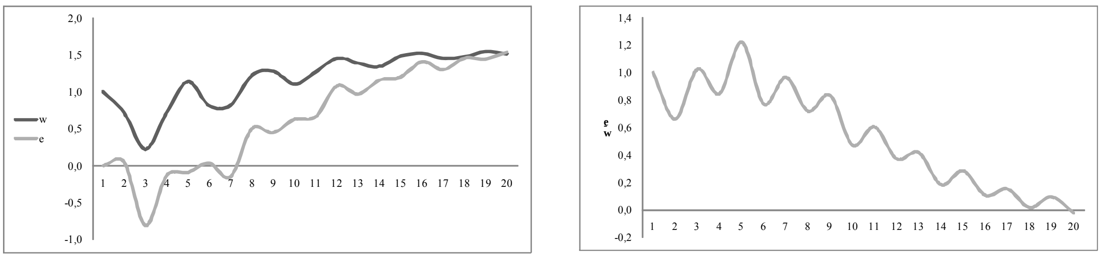
	Regime 1 [1991:4-2008:4]		Regime 2 [1987:2-1991:3]	
	50%	90%	50%	90%
w_t	6.25	20.76	3.20	10.62
e_t	-	-	2.71	8.99

Having estimated the model, it is then possible to apply the impulse-response analysis. The end of this quantitative analysis is to document empirically the likely response of the self-employed firm size to a positive paid-employment shock holding the other variables constant.

Figure 1. Response of log E, log W and log *sfs* to a positive paid-employment shock



Regime: 1991:3-2008:4



Regime:1987:2-1991:3

Figure 1 plots the estimated response to a one-unit increase in paid employment and employers to a positive employment shock in the previous period. As we can see, in the past, *i.e.*, in regime 2, for high *sfs*, when a shock occurred the response of both variables (paid employees and employers) would be positive, whereas from 1991:3 – when the *sfs* is below the estimated threshold parameter- paid employment should decrease in order to close the *sfs* gap created by an equilibrium-distorting shock.

Hence, when the *sfs* is below the estimated threshold parameter as consequence of an adverse shock (adverse demand/supply shock or unfavorable changes in the labour market legislation), a decrease in paid-employment ought to occur in order to restore the long-run equilibrium *sfs*. By contrast, in the past, this kind of shock produced downwards pressures for both employers and employees. In sum, it seems that the use of the contracting out and temporary contracts to circumvent the EPL better accommodate the adjustment to optimal sizes to adapt to changing economic conditions. As a result, higher volatility in the paid employment occurs.

4.4. Is the self–employed firm size stable?

In the previous section, we have concluded that the relationship between employers and paid employment depends on the *sfs*. To look for robustness, we might check whether the *sfs* is stable over time. To this end, we follow the methodology proposed by Bai and Perron. The procedure for detecting structural breaks, suggested by Bai and Perron (1998, 2003a, 2003b), can be described as follows. [24] First, we calculate the *UDmax* and *WDmax* statistics. These are double maximum tests, where the null hypothesis of no structural breaks is tested against the alternative of an unknown number of breaks. These tests are used to determine if at least one structural break is present in the Spanish *sfs* series, in logs. In addition, the $SupF_t(\ell)$ are a series of *Wald* tests for the hypothesis of 0 breaks vs. *l* breaks. In our case, the maximum number of breaks ℓ is chosen to be 5. If these tests show evidence of at least one structural break, then the number of breaks can be determined by the $SupF_t(\ell+1/\ell)$. If this test is significant at the 1 percent level, then *l+1* breaks are chosen. In Table 7, the *UDmax*, *WDmax* and the $SupF_t(\ell)$ statistics suggest the presence of at least one break in the *sfs* variable. In light of this, the number of breaks can be chosen by the $SupF_t(\ell+1/\ell)$ test. For the Spanish *sfs*, the $SupF_t(\ell+1/\ell)$ test is not significant for any ℓ, *i.e.*, no more than 1 break exists in the series.

[24] We apply the Bai-Perron test with only a constant as regressor (i.e. z_t={1}) and account for potential serial correlation via non-parametric adjustments. When implementing the procedure, we used a trimming ε_t=0.15, hence each segment has at least 13 observations. We also allowed serial correlation in the errors and different variances of the residuals across segments.

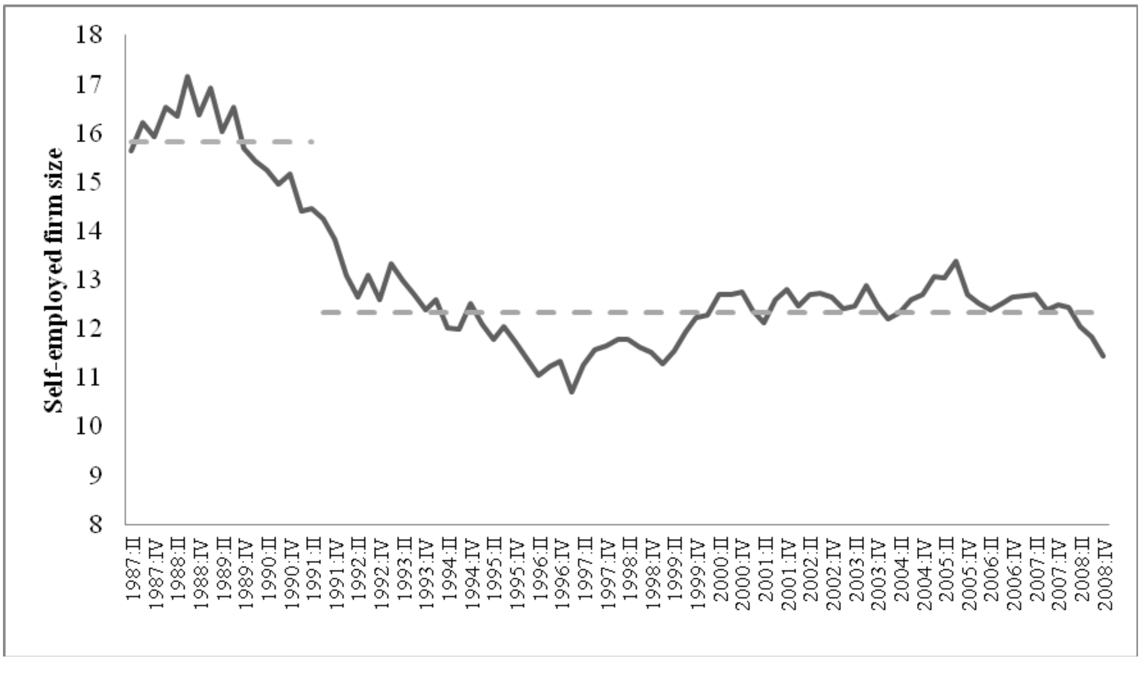
Table 7
Bai-Perron tests of multiple structural changes in the logarithm of average firm size, 1987:2-2008:4

Specifications						
$y_t=ln(W_t/E_t)$	$z_t=\{1\}$	$q=1$	$p=0$	$h=15$	$M=5$	
UDmax	*WDmax*	$SupF_t(1)$	$SupF_t(2)$	$SupF_t(3)$	$SupF_t(4)$	$SupF_t(5)$
51.47^{***}	83.23^{***}	19.65^{***}	3.27	51.47^{***}	38.66^{***}	30.75^{***}
$SupF_t(2/1)$	$SupF_t(3/2)$	$SupF_t(4/3)$	$SupF_t(5/4)$			
2.34	14.00	0.14	0.00			
Break dates estimates						
T_1	1991:3	[1991:1-1993:2]				

Notes: y_t, z_t, q, p, h, and M denote the dependent variable, the explanatory variable allowed to change, the number of regressors, the number of corrections included in the variance-covariance matrix, the minimum number of observations in each segment, and the maximum number of breaks, respectively.
*,**,and *** denote significance at the 10%, 5% and 1% levels, respectively.
The critical values are taken from Bai and Perron (1998), Tables 1 and 2; and from Bai and Perron (2001), Tables 1 and 2. The number of breaks (in our case, one) has been determined according to the sequential procedure of Bai and Perron (1998), at the 1% size for the sequential test.
90% confidence intervals for T_1 in brackets.

Figure 2 presents a graph of the average firm size in Spain from 1987:2 to 2008:4. Our interest is the presence of an abrupt structural change in the mean of the series. The estimated break date is the observation 1991:3, and the development of the *sfs* and its break date are illustrated in Figure 2. It is noteworthy that the point estimate of the break date in the *sfs* is remarkably similar to the estimated threshold. The break date corresponding to the threshold confirms our previous finding as regards the presence of structural breaks in the *sfs*.

Figure 2. Spanish *sfs* development and structural breaks



5. Conclusions

This paper has presented empirical evidence in support of the existence of a change in the way that employers hire workers and the consequences on the Spanish self-employed firm size. Our estimation results reject the null hypothesis of linear cointegration in favour of a two-regime threshold cointegration model, that is, in favour of a time-varying relationship. Our results suggest that 1991:3 marked a break point not only in the *sfs* but also in the way that Spanish employers have responded to changing economic conditions since then.

In particular, our empirical estimates can be explained by means of the two following conjectures. We find evidence of the existence of certain stability in the *sfs* over the long-run. This phenomenon may be the result of the Spanish Labour Market regulation. In fact, Spain has one of the strictest regulatory frameworks among the OECD countries (Faggio and Nickell, 2006). In this context, Spanish entrepreneurs are conscious that in times of crisis it would be very difficult to adjust. The rational behaviour is a conservative position with respect to the contracting of personnel when a positive shock occurs, maintaining a normal demand, and contracting out, even in economic booms. Usually, faced with a positive demand shock, Spanish entrepreneurs have resorted to outsourcing or externalization instead of hiring new employees. This behaviour may be the foundation of the stability showed by the Spanish *sfs*.

However, the new iterations of temporary contracts (without dismissal costs) introduced by Spanish governments in the nineties changed this situation, leading to a change in the mean of the Spanish *sfs* and the adoption of a new strategy by Spanish employers.

The second empirical finding of this paper, the response to a temporary shock in the *sfs*, can also be explained using the same arguments. The introduction of these new flexible formulas led new ways to restore the *sfs*, not by means of simultaneous adjustments by employers and employees but instead by the hiring or firing of workers. As a result, this phenomenon can provide a possible explanation of the extreme volatility recently exhibited by paid employment in Spain.

The existing policy debate over the need to again reform the labor market seems to have ignored this aspect of the problem, *i.e.*, that Spanish entrepreneurs have merely chosen to circumvent the employment protection laws. If this hypothesis is true, the

Spanish unemployment problem should have deeper roots related to the ability of Spanish entrepreneurs to maximize their profits.

Another policy implication of our work is that a larger number of entrepreneurs is not necessarily better if it is the result of contracting out or of the development of new ways of dependent self-employment.

Perhaps the most important policy implication arising out of our analysis is simply to emphasize that the low average contribution to employment of Spanish self-employed workers is the result of a rational response to the change-averse disposition of unions and government. There is therefore a need for Spanish authorities to critically evaluate the correctness of policy choices as regards both the promotion of self employment and employment protection legislation.

The alternative to the current situation is to attempt to implement institutional changes which somehow disincentivize contracting out and the hiring of temporary workers and reduce the 'false' flows from paid employment to self-employment.

We are aware of certain limitations in our analysis. First, the sfs could develop in different sectors of economy in different ways. Second, we must further research the relationship between the self-employed firm size and other factors. This should be a key issue for further research.

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Appendix A: Figures

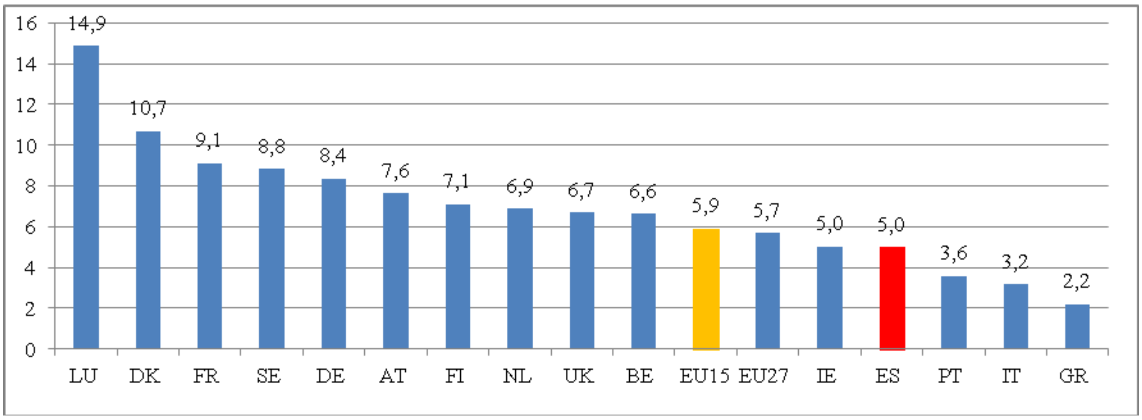

Figure A1: Self-employed firm size in EU-15 countries, 2008. Ratio paid-employees/self-employed. Source: Labour Force Survey, Eurostat.

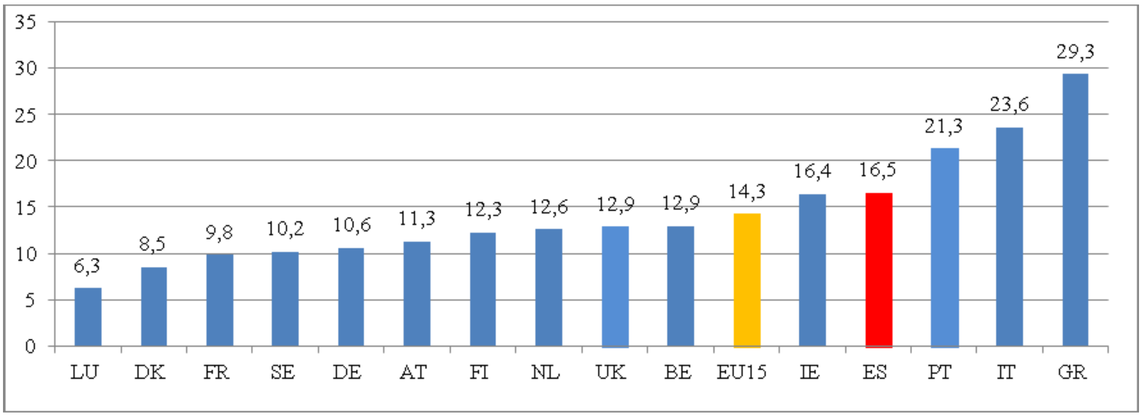

Figure A2: Self-employment rates in EU-15 countries, 2008. Source: Labour Force Survey, Eurostat.

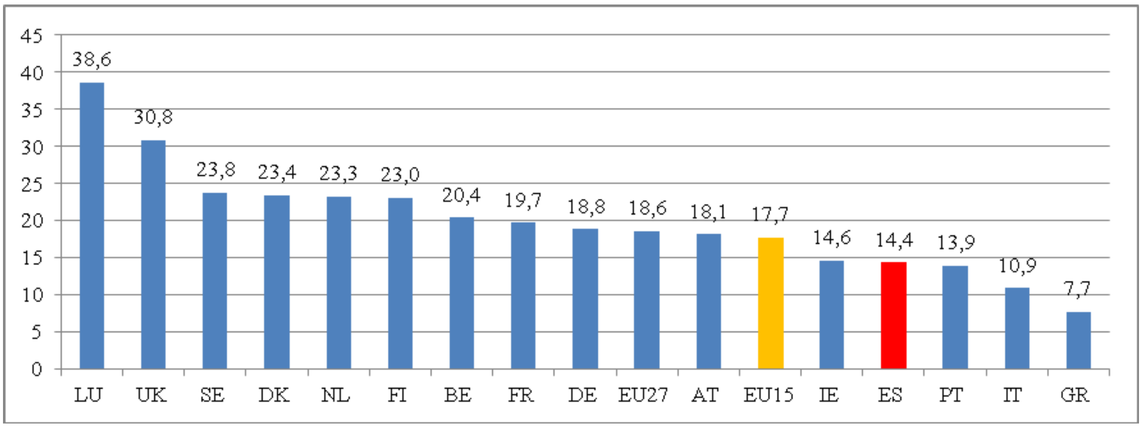

Figure A3: Self-employed firm size in EU-15 countries, 2008. Ratio paid-employees/employers. Source: Labour Force Survey, Eurostat.

Appendix B: Threshold VECM estimates with non-fixed beta

Table C1
Threshold VECM Estimates, non-fixed beta (Hansen-Seo, 2002)

Threshold	Regime 1 $\left(w_{t-1} - 0.969e_{t-1}\right) \leq 2.86$ $sfs \leq 17.46$		Regime 2 $\left(w_{t-1} - 0.969e_{t-1}\right) > 2.86$ $sfs > 17.46$	
Variable	Δw_t	Δe_t	Δw_t	Δe_t
C	0.262***	-0.042	-0.580***	-0.710*
	(0.084)	(0.163)	(0.144)	(0.377)
Δw_{t-1}	0.420***	-0.083	-0.484***	-0.015
	(0.135)	(0.323)	(0.175)	(0.249)
Δw_{t-2}	-0.189**	0.036	-0.768***	-1.069***
	(0.082)	(0.220)	(0.154)	(0.425)
Δe_{t-1}	-0.032	-0.157	0.187**	-0.015
	(0.068)	(0.104)	(0.079)	(0.249)
Δe_{t-2}	-0.079	-0.097	0.008	0.662***
	(0.070)	(0.145)	(0.071)	(0.187)
α	-0.094***	0.021	0.204***	0.248**
	(0.031)	(0.059)	(0.049)	(0.127)
Observations percentage:	82.14 %		17.86 %	

Notes: Standard errors in brackets.
*, **,*** Significant at the 10%, 5% and 1% levels, respectively.