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The dynamics of entrepreneurship: hysteresis, business cycles and government policy

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Abstract This article estimates two unobserved components models to explore the macrodynamics of entrepreneurship in Spain and the USA. We ask whether entrepreneurship exhibits *hysteresis*, defined as a macrodynamic structure in which the cyclical component of entrepreneurship has persistent effects on the natural rate of entrepreneurship. We find evidence of hysteresis in Spain, but not the USA, while in Spain business cycle output variations significantly affect future rates of entrepreneurship. The article discusses implications of the findings for the design of entrepreneurship policies.

Keywords Hysteresis · Unobserved components model · Time series models · Business cycles · Self-employment · Entrepreneurship

JEL Classification C32 · J23 · M13

1 Introduction

As national economies continue to wrestle with the forces of globalization, and large companies proceed with outsourcing and downsizing strategies, efforts to find

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alternative sources of economic growth are intensifying. For many years, governments around the world have regarded entrepreneurship as a promising candidate in this respect. Growing evidence shows that entrepreneurs create disproportionate numbers of innovations and jobs (Acs and Audretsch 1990; Audretsch 2003; Haltiwanger 2006). Entrepreneurship has also been linked with faster rates of economic growth (Audretsch and Keilbach 2004; van Stel et al. 2005).

Many governments have responded to these forces by devising and implementing portfolios of policies to promote entrepreneurship. These policies include loan guarantee schemes; technology-transfer and innovation programs; employment assistance programs; and subsidized provision of business advice and assistance to small firms (Parker 2009). Loan guarantee schemes insure banks' loans to entrepreneurs; high rates of business failure mean that these schemes typically run at a loss (Parker 2009, chapter 16). Innovation policies include direct subsidies to innovators; favorable tax treatment for private sector R&D expenditures; and the provision of seed funds for innovation (Lerner 1999). Employment assistance programs subsidize welfare recipients to leave the unemployment register by starting new ventures (Bendick and Egan 1987). Taken together, these interventions often impose sizeable costs on the taxpayer. For example, the Inno-Empresa program to promote entrepreneurship in Spain has an annual budget of €735 million, while the US Small Business Innovation Research Program offers about $2.5 billion per year in awards (OECD 2010).

Given these costs, the lack of robust evidence associating these policies with expanded levels of entrepreneurship is particularly striking. Part of the difficulty of evaluating entrepreneurship policies is that they may have very long run effects. For example, regional and national data suggest that some entrepreneurship outcomes, especially employment creation and venture growth, can take a decade or more to play out (Fritsch and Mueller 2008; Carree and Thurik 2008). These long-run effects are not accurately captured by conventional evaluations, which are usually performed a few years after the policies are implemented, and so capture only short-term impacts (Hart 2003). An important question therefore concerns the durability of shocks to entrepreneurship, whether these are "policy" shocks (i.e., derived from sudden changes to government policy) or "economic" shocks (e.g., derived from sudden changes to technology).

At the heart of this question is whether entrepreneurship evolves as a trend-stationary or as a non-stationary time-series process. If entrepreneurship is trend-stationary, economic and policy shocks can be regarded as transitory from an aggregate perspective: the rate of entrepreneurship eventually reverts to its underlying, long-run ("natural") rate. Granted, this "natural rate" might also shift over time; but then one would expect entrepreneurship to be stationary once structural breaks are allowed for. For instance, one can conceive of a policy which suddenly liberalizes entry into previously closed markets, and which thereby creates an upward shift in the amount of entrepreneurship in the economy. If the rate of entrepreneurship is stationary, the shock dies away once the policy has been implemented, and entrepreneurship settles down at its new, higher, rate. If on the other hand the rate of entrepreneurship is non-stationary, such shocks can have permanent effects. For example, the one-off market entry liberalization policy considered above increases the rate of entrepreneurship in a way which affects all subsequent rates of entrepreneurship. A necessary (but not sufficient) condition for this to occur is a unit root in the rate of entrepreneurship.

In a time-series context, hysteresis can be defined and measured in various ways. A popular approach in the empirical literature simply equates hysteresis with the existence of a unit root in a variable (see, Røed 1997, for a survey). An alternative approach proposed by Jaeger and Parkinson (1990, 1994) posits a more demanding criterion: hysteresis exists if shocks (such as the one-off policy change discussed above) affect the natural rate of a variable, which itself follows a unit root process. In which case, temporary shocks have permanent effects while the business cycle does not evolve independently of the natural rate; it then follows that a unit root is a necessary but not a sufficient condition for hysteresis. In this article, we adopt Jaeger and Parkinson (1990, 1994) definition of hysteresis to conduct a searching test and to explore whether entrepreneurship covaries systematically with the business cycle.

To test for hysteresis in this way, we follow Jaeger and Parkinson (1990, 1994) and decompose entrepreneurship into two unobservable components: a non-stationary "natural rate" component, and a stationary "cyclical" component. These components can be estimated by maximum likelihood using the Kalman filter. Although Jaeger and Parkinson's approach has been applied extensively in the literature on unemployment (see Assarson and Janson 1998; Karamé 1999; Salemi 1999; Pérez-Alonso and Di Sanzo 2011; Logeay and Tober 2005), to the best of our knowledge its application to entrepreneurship is novel.

The goal of this article is to explore whether aggregate rates of entrepreneurship exhibit persistence or hysteresis. We do so using quarterly time-series data from Spain and the USA. These two countries have experienced some interesting differences over the last three or four decades. For example, self-employment rates in Spain tend to be relatively high by OECD standards and are on a rising trajectory, having increased from 11.6% in 1972 to 12.3% by 1989 and further to 13.3% by 2006 (Wennekers et al. 2008). In contrast, US self-employment rates were 8.2% in 1972, increasing to 10.8% by 1989 before falling back to 10.1% by 2006.

The two countries have also had different experiences with unemployment over this period, with Spain typically having higher, more persistent and more variable unemployment rates than the USA (Eurostat 2009a). Institutional changes in Spain seem to have played a major role in this regard, with a transition from dictatorship to democracy in 1975 triggering a set of deep economic reforms designed to increase the country's competitiveness and openness. This was associated with rising unemployment rates, from under 5 per cent in the mid-1970s to over 20% by the mid-1980s. This episode was followed by Spain's entrance into the European Union (EU) in 1986, which was followed by reductions in the unemployment rate to 16%. Although the recession of the early 1990s returned Spanish unemployment rates to over 20% again, a long period of ongoing fiscal and regulatory reform and active labor market programs, accompanied by a prolonged global economic upswing, progressively reduced unemployment rates to below 8% by 2007.

According to Eurostat (2009b, Table 3) after Germany, Spain has the most extensive set of active labor market programs devoted to encouraging start-ups of any country in the EU. Yet it remains unclear to what extent these programs, as well as its regulatory environment and welfare system, impact entrepreneurship and hysteresis in Spain relative to the more laissez faire USA (Di Tella and MacCulloch 2006; Raurich et al. 2006). On the one hand, less regulated markets might be associated with lower entry barriers, higher rates of entrepreneurship and also higher rates of exit (Fonseca et al. 2001). On the other hand, protective institutions such as unemployment benefit entitlements, employment protection laws, and trade unions might reduce personal risks entailed with entrepreneurial entry, while having limited effects on national unemployment rates (Baker et al. 2005). The implications for hysteresis in entrepreneurship in both cases are unclear, motivating an empirical analysis of this issue.

A further motivation for our empirical analysis is that hysteresis implies that entrepreneurship policies might be more powerful than has been thought hitherto. That is because any increase in entrepreneurship brought about these policies will be incorporated into all future levels of entrepreneurial activity if hysteresis exists. Furthermore, business cycles could have important effects on the real economy, by affecting the future trajectory of an economy's natural rate of entrepreneurship.

This article has the following structure. Section 2 discusses in greater detail theory and evidence about persistence, hysteresis and business cycles in entrepreneurship. Section 3 describes the data and the estimation methodology. Section 4 presents and discusses the results and performs a robustness check on the specification of the model. Section 5 concludes with a discussion of policy implications and some promising avenues for future research.

2 Hysteresis, persistence, business cycles, and entrepreneurship

Rates of entrepreneurship vary dramatically between countries but exhibit a fairly high degree of temporal stability (Parker and Robson 2004). Individual-level panel data reveal that the best predictor of someone being self-employed in the next period is whether they are self-employed in the current period (Henley 2004). This "state-dependence" property appears to aggregate up to the regional level. For example, Fritsch and Mueller (2007) explain more than one-half of the variance in German regional start-up rates in terms of regional start-up rates from 15 years earlier. The same property also holds at the national level, with several studies being unable to reject the null hypothesis of a unit root in self-employment rates (Parker 1996; Cowling and Mitchell 1997; Parker and Robson 2004; Bruce and Mohsin 2006).

What might explain these findings? At the individual level, there could be non-pecuniary costs of switching occupation, such as the sudden loss of a pleasant compensating differential, disruption to an accustomed lifestyle, or a stigma from failure (Landier 2004). Alternatively, switching costs could be economic in nature involving, for example, lost sector-specific experience, costs of raising start-up capital (if entering entrepreneurship), or re-training costs (if entering paid-employment). Switching costs might also relate to exit barriers caused by incurring sunk costs of capital with limited resale value; prior commitments to customers; or a desire by entrepreneurs to avoid sending an adverse signal of ability by abandoning their ventures (Boot 1992). In a different vein, Dixit (1989) and Dixit and Rob (1994) show that risk together with sunk costs can give agents an option value of waiting before switching occupation. This reduces the total amount of entry and exit that occurs—as conditions have to become very bad before entrepreneurs close their business and relinquish their sunk costs, or very favorable before they are willing to incur the risk of jeopardizing their assets by entering the market. Risk generates an "option value" of remaining in the present occupation and deferring a costly switch. Only when average incomes in entrepreneurship reach some upper "trigger point" will people become entrepreneurs. And they will only leave entrepreneurship in the presence of adjustment costs if incomes drop to some lower trigger point. Between these two trigger points individuals remain in their current occupation. Dixit and Rob (1994) explicitly refer to this inertia in occupational choice as "hysteresis".

At the more aggregated level, theoretical models of multiple entrepreneurship equilibria can explain why ostensibly similar regions and countries exhibit pronounced and enduring differences in entrepreneurship. Thus, Landier (2004) studies a model in which serial entrepreneurs possess private information about their abilities which cannot be credibly revealed to banks. High-quality serial entrepreneurship is deterred in economies where the equilibrium cost of capital is high. The cost of capital is high precisely because there is little or no high-quality serial entrepreneurship. But high quality serial entrepreneurship becomes privately worthwhile in economies where the equilibrium cost of capital is low—which in turn justifies the low cost of capital. Another multiple equilibrium model, by Parker (2005), explains why different geographical areas can possess persistently different rates of entrepreneurship based on self-perpetuating human capital choices within regions which affect payoffs in entrepreneurship and in paid-employment, locking different occupational choice structures into place as stable equilibria.

We turn now to the relationship between business cycles and entrepreneurship. In principle, entrepreneurship could evolve either pro- or anti-cyclically, depending on the balance of forces at work in the private sector of the economy. Several theories posit a pro-cyclical relationship. For example, Rampini (2004) argues that favorable shocks to the economy increase productivity and wealth, making agents more willing to bear risk (via decreasing absolute risk aversion) and become entrepreneurs. And, anticipating greater returns in favorable states, entrepreneurs also supply higher levels of effort, reducing moral hazard problems and making lenders more willing to fund risky investment projects. When shocks are unfavorable, the opposite process occurs: wealth, investment, and entrepreneurship all decline. Shleifer (1986) suggests another mechanism, based on the idea of "implementation cycles" in innovation. When entrepreneurs anticipate that imitators will erode profits from costly innovation, they optimize by delaying the release (implementation) of innovations until economic conditions are favorable. The outcome is pro-cyclicality into the temporal evolution of entrepreneurship and innovation. This effect can be exacerbated when there a dynamic externality inherent in innovation (Barlevy 2007). Thus radical innovations increase economic activity directly—and frequently indirectly create opportunities for other, subsequent innovations, further increasing opportunities for entrepreneurship and greater economic activity. Because entrepreneurs do not internalize this dynamic externality when making their decisions to innovate and invest, the result is volatility and pro-cyclicality of entrepreneurship, innovation and economic growth.

Yet other theories predict counter-cyclical entrepreneurship. Francois and Lloyd-Ellis (2003) identify a limitation in Shleifer (1986) analysis: if entrepreneurs can store their output, they would do best innovating and producing when costs are low (during recessions) and selling when demand is high (during booms). This separation of production and sales gives entrepreneurs incentives to enter during recessions. A counter-cyclical effect of this sort is enhanced when entrants can exploit new technologies and displace incumbents (Caballero and Hammour 1994). Displacement and the consequent reallocation of resources occurs most frequently during recessions: Caballero and Hammour point to the displacement of US artisan automobile producers by new mass producers during the Great Depression as an example of this phenomenon. Caballero and Hammour refer to the role of entrepreneurs introducing superior new technology and displacing inefficient incumbents as the "cleansing effect of recessions". Counter-cyclicality is also consistent with the emergence of worker co-operatives and other "marginal" enterprises in recessions, which dissolve in economic recoveries when conventional employment opportunities become more readily available (Ben-Ner 1988; Pérotin 2006).

A further argument for counter-cyclicality of entrepreneurship relates to monetary policy, since the cost of capital tends to increase in booms and decrease in recessions, inducing exits in the former state and entries in the latter. Of course, aggregate market demand is also higher in booms and lower in recessions, which could offset changes in the cost of capital in terms of occupational choice. The entrepreneurship literature has referred to these conflicting forces in terms of "recession push" and "prosperity pull" effects (Parker 2009, Chapter 4).

The available evidence broadly suggests that venture formation rates and individual transitions into entrepreneurship are higher on average in good economic times and lower on average in bad ones (Audretsch and Acs 1994; Grant 1996; Carrasco 1999). However, this evidence is rather informal in nature. It is based on estimates of the sign of time dummies in individual-level studies of occupational choice rather than being derived from careful analyses of time-series data. It will therefore be interesting to see whether the results obtained in this article, derived using a dynamic time-series estimation methodology, bear out these suggestive findings.

Finally, we believe that previous entrepreneurship research has overlooked an important distinction between different types of entrepreneurs. Entrepreneurs who hire external labor ("employers") belong to a distinct group which could exhibit different cyclical behavior compared with entrepreneurs who work on their own ("own-account entrepreneurs").[1] Both types of entrepreneur are likely to benefit from higher demand (growth in national income). But employers who run larger ventures and so benefit from economies of scale are likely to gain the most from demand growth (Klepper 1996). These entrepreneurs can scale up production and expand employment, bidding up wages which draw relatively low-value own-account entrepreneurs out of entrepreneurship and into paid-employment (Lucas 1978). In which case, one might expect the number of employer entrepreneurs to increase relative to the number of own-account entrepreneurs during booms, making cyclical effects positive for

[1] To date the differences between these two groups have not been considered to any great extent in the literature (some exceptions are Carroll et al. 2000; Burke et al. 2002; Cowling et al. 2004).

employer entrepreneurs and negative for own-account entrepreneurs. And to the extent that more favorable economic conditions improve opportunities for some own-account entrepreneurs as well, we might expect the latter to start hiring labor (Cowling et al. 2004), in which case they switch from own-account to employer status, and reinforce the positive cyclical effects for employers and the negative cyclical effects for the own-account group. Our empirical estimates below will shed light on these conjectures.

3 Data and methodology

3.1 Data

The first task is to measure entrepreneurship. Entrepreneurship is a multifaceted concept, which encompasses a range of roles and activities, including risk-bearing, co-ordination, arbitrage, and Schumpeterian innovation (Iversen et al. 2008). Any single measure of entrepreneurship is therefore unlikely to do justice to all of these different facets. In cross-country comparisons, by far the most common measure used in practice is self-employment rates (Iversen et al. 2008; Parker 2009). This chiefly reflects the widespread availability of aggregate self-employment data for a range of countries. As noted by entrepreneurship scholars, the self-employment definition has the merit of inclusiveness and convenience. And by being residual claimants of their own ventures, the self-employed correspond to the risk-bearing arbitrageur entrepreneur emphasized in the writings of of Knight, Say, and Kirzner (Iversen et al. 2008; O‘kean and Menudo 2008).

However, self-employment is not a perfect measure of entrepreneurship, since it includes numerous “casual” businesses as well as long-established enterprises. As such, the self-employment rate may represent long-standing industrial and institutional structures within a country rather than the dynamism of an entrepreneurial economy. Self-employment data also typically under-sample Schumpeterian innovative entrepreneurs relative to “replicative” or “me too” businesses (Audretsch 2002; Baumol et al. 2009). This consideration might help explain, for example, why Spain has higher self-employment rates than the USA, despite the fact that the USA is widely regarded as the world’s leader in entrepreneurship (Iversen et al. 2008). Because we were unable to find an alternative (or additional) measure of entrepreneurship, these limitations should be borne in mind in the analysis below. In particular, we have to acknowledge the possibility that any hysteresis in self-employment rates may actually reflect the presence of country-specific institutions rather than the explanations discussed in Sect. 2 above.

Our empirical analysis uses seasonally adjusted[2] as well as seasonally unadjusted quarterly data on self-employment rates, for the US and Spain. Many previous authors exclude the self-employed in the agricultural sector because this sector is structurally different from the rest of the economy.[3] In what follows, we will present results both

[2] The seasonal adjustment procedure is the U.S. Census Bureau’s X12 ARIMA process.

[3] Previous authors have pointed out that self-employment in agriculture is likely to be heavily influenced by historically and culturally determined traditions of family ownership and factors (such as a high proportion

with and without the agricultural sector included. The self-employment rate, (S_t), is defined as the share of employed people that is self-employed in non-agricultural activities. Rates of employer self-employment (E_t) and own-account self-employment (OA_t) are defined as the number of employers and own-account workers respectively, divided by total employment.

The US self-employment data are seasonally adjusted quarterly observations drawn from the Current Population Survey (CPS, US Bureau of Labor Statistics). The Spanish self-employment data are seasonally adjusted quarterly observations drawn from the Labor Force Survey (EPA, Spanish National Statistics Institute). Owing to Spanish data limitations, both samples start in 1987(II) and conclude in 2008(IV). A longer time series is available when agricultural self-employment is included.[4]

Finally, real GDP is denoted by Y_t. Data on Spanish real GDP are taken from the Quarterly National Accounts database while data on US GDP are taken from the US Department of Commerce. These data are seasonally adjusted and are expressed in billions of chained 2005 US dollars.

3.2 Econometric methodology

Several macroeconomic studies equate hysteresis in a time series with a unit root process.[5] Others argue that hysteresis arises when changes to the cyclical component of a time series, S_t^{C}, induce permanent changes in the "natural rate" of the series, S_t^{N}. This is different to a unit root process. To comprehend the different estimation strategies these approaches call for, decompose the series S_t into the sum of its two (unobservable) components: the non-stationary natural rate component, S_t^{N}, and the stationary cyclical component, S_t^{C}:

$$S_t = S_t^{\mathrm{N}} + S_t^{\mathrm{C}} \tag{1}$$

Now define the natural rate component as a random walk plus a term capturing a possible hysteresis effect:

$$S_t^{\mathrm{N}} = S_{t-1}^{\mathrm{N}} + \beta S_{t-1}^{\mathrm{C}} + \varepsilon_t^{\mathrm{N}} \tag{2}$$

where the β coefficient measures, in percentage points, how much the natural rate increases if the economy experiences a cyclical self-employment rate increase of 1%.

Footnote 3 continued
of unpaid family workers) other than those that influence other types of self-employment (Iversen et al. 2008). Note that the self-employed are categorized differently by the American CPS compared with the Spanish EPA—in a way which increases the share of workers classified as self-employed in Spain relative to the US. In the US, independent owner-managers and directors of *incorporated* enterprises are classified as employees, while in Spain they are classified as employers. In addition, the Spanish data allow the researcher to distinguish between own-account workers and employers, whereas they cannot be separated in the US case.

[4] The second sample spans 1980(I) through 2008(IV).

[5] See Blanchard and Summers (1986). Layard et al. (1991) popularized the term "pure" hysteresis for describing the presence of a unit root in time series.

Evidently a unit root in the self-employment rate S_t is a necessary but not sufficient condition for the existence of hysteresis since a unit root could be generated by an accumulation of shocks to the natural rate S_t^{N}while at the same time $\beta = 0$ (Røed 1997). In contrast, there is hysteresis if $\beta > 0$.

The specification of the model is completed by writing the cyclical component of the self-employment rate as a stationary second-order autoregressive process[6]:

$$S_t^{\mathrm{C}} = \varphi_1 S_{t-1}^{\mathrm{C}} + \varphi_2 S_{t-2}^{\mathrm{C}} + \varepsilon_t^{\mathrm{C}} \tag{3}$$

where ϕ_1 and ϕ_2 provide a measure of the periodicity of the cyclical component[7]. The random shocks $\varepsilon_t^{\mathrm{N}}$ and $\varepsilon_t^{\mathrm{C}}$ are assumed to be mean-zero draws from the normal distribution with variance–covariance matrix Ω; the state-space form[8] of the model can be written as

$$S_t = \begin{pmatrix} 1 & 1 & 0 \end{pmatrix} \begin{pmatrix} S_t^{\mathrm{N}} \\ S_t^{\mathrm{C}} \\ S_{t-1}^{\mathrm{C}} \end{pmatrix} \tag{4}$$

$$\begin{pmatrix} S_t^{\mathrm{N}} \\ S_t^{\mathrm{C}} \\ S_{t-1}^{\mathrm{C}} \end{pmatrix} = \begin{pmatrix} 1 & \beta & 0 \\ 0 & \varphi_1 & \varphi_2 \\ 0 & 1 & 0 \end{pmatrix} \begin{pmatrix} S_{t-1}^{\mathrm{N}} \\ S_{t-1}^{\mathrm{C}} \\ S_{t-2}^{\mathrm{C}} \end{pmatrix} + \begin{pmatrix} \varepsilon_t^{\mathrm{N}} \\ \varepsilon_t^{\mathrm{C}} \\ 0 \end{pmatrix} \tag{5}$$

$$\Omega = \begin{pmatrix} \sigma_N^2 & 0 & 0 \\ 0 & \sigma_C^2 & 0 \\ 0 & 0 & 0 \end{pmatrix} \tag{6}$$

To summarize, hysteresis is inferred if the coefficient β is significantly different from zero. The coefficients of the model (4)–(6) are estimated by maximum likelihood using a Kalman filter.

A non-linear version of this model (4) through (6) can also be estimated, to take account of the possibility that entrepreneurship rates displays asymmetries in adjustment dynamics in response to positive and negative shocks. Relaxing the linearity assumption may allow a better estimation of hysteresis if it exists. When we talk about "positive" or "negative" shocks, we do so relative to some threshold level of self-employment rate, τ(where $\tau$is not necessarily zero). To explore whether asymmetries exist, we estimate a non-linear version of the unobserved components model by allowing past cyclical self-employment to have a different impact on the natural rate, which depends on the regime of the economy. Specifically, we replace the state-space

[6] The assumption of a purely autoregressive process for the self-employment cycle can be relaxed in favor of more general (and possibly more parsimonious) autoregressive moving-average specifications. In the present application, an AR(2) fits the data best according to AIC comparisons. Full results are available from the authors on request.

[7] Stationarity implies that the roots of the polynomial equation $1 - \phi_1 L - \phi_2 L^2 = 0$; where L is the lag operator, should lie outside the unit circle.

[8] Proietti (2004) states that this model is identified without any additional identification equation being required.

equation (5) with the Threshold Auto Regressive (TAR) specification

$$\begin{pmatrix} S_t^{\mathrm{N}} \\ S_t^{\mathrm{C}} \\ S_{t-1}^{\mathrm{C}} \end{pmatrix} = \begin{pmatrix} 1 & 0 & 0 \\ 0 & \varphi_1 & \varphi_2 \\ 0 & 1 & 0 \end{pmatrix} \begin{pmatrix} S_{t-1}^{\mathrm{N}} \\ S_{t-1}^{\mathrm{C}} \\ S_{t-2}^{\mathrm{C}} \end{pmatrix} + \begin{pmatrix} \beta^+ \\ 0 \\ 0 \end{pmatrix} I_t^+ S_{t-1}^C + \begin{pmatrix} \beta^- \\ 0 \\ 0 \end{pmatrix} I_t^- S_{t-1}^C + \begin{pmatrix} \varepsilon_t^{\mathrm{N}} \\ \varepsilon_t^{\mathrm{C}} \\ 0 \end{pmatrix} \tag{7}$$

where I_t^+ and I_t^- are the Heaviside indicator functions such that:

$$I_t^+ = \begin{cases} 1 & \text{if} \quad S_{t-1}^{\mathrm{C}} \geq \tau \\ 0 & \text{if} \quad S_{t-1}^{\mathrm{C}} < \tau \end{cases}$$
$$I_t^- = \begin{cases} 1 & \text{if} \quad S_{t-1}^{\mathrm{C}} < \tau \\ 0 & \text{if} \quad S_{t-1}^{\mathrm{C}} \geq \tau \end{cases}$$

This model can be estimated via maximum likelihood using the Kalman filter, where τ is unknown so and it is estimated along with the other parameters of the model β^+ and β^-. In this context a test for asymmetry becomes a test for linearity, i.e., a test for a single regime against the alternative of two regimes. The null hypothesis we are interested in is $H_0 : \beta^+ = \beta^-$ vs. $H_0 : \beta^+ \neq \beta^-$. If we reject the null of linearity, there is evidence for the presence of a type of nonlinear hysteresis in the self-employment rate, i.e., with cyclical shocks being propagated asymmetrically to the natural rate. Given our model, the asymptotic distribution of conventional test statistics is not χ^2.[9] To circumvent this problem we follow Pérez-Alonso and Di Sanzo (2011) who suggest using bootstrap methods to approximate the sampling distribution of the test statistic.

Because the relationship between business cycles and entrepreneurship is also of interest we also estimate a bivariate correlated unobserved components model to investigate the interplay between output and the self-employment rate. We use Sinclair's (2009) extension of Clark's (1989) model which allows for possible correlation between the components of the covariance matrix. This model can be used to decompose self-employment and output into their cyclical and natural components:

$$S_t = S_t^{\mathrm{N}} + S_t^{\mathrm{C}} \tag{8}$$
$$Y_t = Y_t^{\mathrm{N}} + Y_t^{\mathrm{C}} \tag{9}$$

Each cyclical component is modeled as an AR(2) process.[10]:

$$S_t^{\mathrm{C}} = \varphi_{1s} S_{t-1}^{\mathrm{C}} + \varphi_{2s} S_{t-2}^{\mathrm{C}} + \varepsilon_{st} \quad \varepsilon_{st} \text{ NID} \left(0, \sigma_{\varepsilon_s}^2\right) \tag{10}$$
$$Y_t^{\mathrm{C}} = \varphi_{1y} Y_{t-1}^{\mathrm{C}} + \varphi_{2y} Y_{t-2}^{\mathrm{C}} + \varepsilon_{yt} \quad \varepsilon_{yt} \text{ NID} \left(0, \sigma_{\varepsilon_y}^2\right) \tag{11}$$

9 See e.g., Hansen (1999) and Lo and Zivot (2001).

10 We could use an AR(p) process for both cyclical components. However, we find that an AR(2) process for the cyclical component fits the data best according to the AIC criterion

Each natural component is assumed to be given by a random walk, although we also allow for a drift in the GDP equation:

$$S_t^{\mathrm{N}} = S_{t-1}^{\mathrm{N}} + \eta_{st} \quad \eta_{st} \ \mathrm{NID}\left(0, \sigma_{\eta_s}^2\right) \tag{12}$$

$$Y_t^{\mathrm{N}} = \mu_t + Y_{t-1}^{\mathrm{N}} + \eta_{yt} \quad \eta_{yt} \ \mathrm{NID}\left(0, \sigma_{\eta_y}^2\right) \tag{13}$$

The state-space form of this model can be estimated using the Kalman Filter with maximum likelihood estimation of the parameters and the cyclical and natural components.[11] We also estimate all the correlations between the unobserved components of the two series. The correlation coefficient between the cyclical component of GDP and self-employment rate reveals pro- or counter-cyclical variation depending on whether the coefficient is positive or negative. At the same time, the correlation between the natural components of GDP and self-employment rate reveals the nature of the long run relationship between these variables.

4 Results

This section presents the results in several stages. First, we test what Jaeger and Parkinson (1990, 1994) have characterized as a necessary but not sufficient condition for hysteresis, namely the existence of a unit root in the self-employment time-series.[12] Because unit root tests are well-known, our discussion will be deliberately brief. Second, we estimate the linear unobserved components model outlined in Sect. 3.2, incorporating a unit root as a maintained hypothesis. This enables hysteresis to be tested directly. The third subsection explores the possibility of asymmetric behavior in adjustment dynamics, by estimating the nonlinear TAR unobserved components model, using the estimation strategy proposed by Pérez-Alonso and Di Sanzo (2011). The relaxation of linearity acts as one important robustness check on the results. Finally the existence of (symmetric) business cycle effects are examined by estimating correlations between cyclical and natural components using Sinclair's (2009) bivariate unobserved component model.

4.1 Unit root tests

In order to test the hypothesis of non-stationarity, we apply the traditional Augmented Dickey–Fuller (ADF) test and a modified version of the Dickey–Fuller and Phillips–Perron tests proposed by Ng and Perron (2001). This comprises a class of modified tests, $\bar{M}$, with GLS de-trending of the data and use of the modified Akaike information Criteria to select the autoregressive truncation lag. Table 1 reports the results of Ng–Perron tests, $\overline{M}Z_{\alpha}^{\mathrm{GLS}}$ and $\overline{M}Z_{t}^{\mathrm{GLS}}$, originally developed in Stock (1999) with GLS

[11] As Sinclair (2009) explains, this model is identified without the imposition of any restrictions on the covariance matrix.

[12] For the model to make sense, a unit root in the self-employment rate is a necessary, and testable, assumption.

de-trending of the data as proposed by Elliot et al. (1996). In addition, Ng–Perron proposed a similar procedure that corrects the problem associated with the standard ADF test, $\overline{M}SB^{\mathrm{GLS}}$ and $\bar{M}PT^{\mathrm{GLS}}$. All test statistics formally examine the unit root null hypothesis against the alternative of stationarity.[13]

The results in Table 1 show that the null hypothesis of non-stationarity cannot be rejected for any series, regardless of the test. However, it is well known that structural breaks in time-series can lead to spurious inferences of a unit root. To deal with this possibility, we employ the Zivot and Andrews (1992) minimum ADF-t(min-t) procedure. The min-t statistics reported in Table 2 show that the null hypothesis of a unit root in the time series still cannot be rejected for either country. This buttresses our conclusion that a unit root exists in the self-employment rates of both Spain and the USA—and additionally, for each type of self-employment in Spain.[14] As noted above, a unit root is a maintained assumption needed to test for Jaeger and Parkinson's notion of hysteresis. We investigate the notion of hysteresis next.

4.2 The linear unobserved component model

Like Tables 1 and 2, Table 3 comprises four blocks of results: with and without agriculture and with and without seasonal adjustment. The first block uses excludes agriculture and is for seasonally adjusted data. The first two columns of Table 3 present the results of estimating (4) through (6) for aggregate self-employment rates in the USA and Spain, respectively. The parameter β is positive in both countries, and significant in Spain but not the USA. The cyclical component of Spanish self-employment is particularly large in magnitude, with parameter estimates suggesting that a 1% increase in the cyclical component of self-employment translates into an increase in the natural rate of Spanish self-employment of as much as 0.283%. The effects of hysteresis for these series are illustrated in Figs. 1 and 2. These figures depict, for Spain and the US, respectively, self-employment rates and estimates of the natural rate and the cyclical component. In Spain, where evidence of pronounced hysteresis has been detected, the natural rate component of self-employment follows quite closely the actual self-employment rate. By contrast, in the US the natural rate of non-agricultural self-employment is rather more stable.

Drawing on our earlier conceptual discussion, we now explore the Spanish data further by decomposing the aggregate self-employment rate into its two constituent parts, employer (E) and own-account (OA) self-employment. We do so in order to determine whether hysteresis is being driven by one or both of these elements. We then apply the unobserved components model (4) through (6) to each of these two constituent self-employment rates separately. The third and fourth columns in the first block of Table 3 contain the results. As can be seen, both components of Spanish self-employment exhibit hysteresis separately. Hysteresis seems to be more pronounced

[13] Other unit root tests allow for the possibility of non-linear behavior: see Papell et al. (2000), León-Ledesma and McAdam (2004), Camarero and Tamarit (2004), and Camarero et al. (2006, 2008).

[14] Similar results are obtained when we allow for multiple structural breaks using the framework of Bai and Perron (1998, 2003a,b). Results are suppressed for brevity but are available from the authors on request.

Table 1 Unit root tests

Dataset	Non-agricultural self-employment Seasonally adjusted data				Non-agricultural self-employment Non-seasonally adjusted data				Self-employment Seasonally adjusted data				Self-employment Non-seasonally adjusted data			
Country	US	Spain			US	Spain			US	Spain			US	Spain		
Status	S	S	E	OA	S	S	E	OA	S	S	E	OA	S	S	E	OA
$\overline{M}Z_{\alpha}^{GLS}$	−0.001	−2.026	0.964	−0.106	−1.818	−2.183	0.781	−0.041	1.541	0.032	−1.706	0.350	1.591	−0.493	0.925	0.289
$\overline{M}Z_{t}^{GLS}$	−0.000	−0.889	0.891	−0.071	0.944	−0.931	0.619	−0.029	0.904	0.020	−0.648	0.229	0.8841	−0.262	0.718	0.185
$\overline{M}SB^{GLS}$	0.395	−0,439	0.924	0.668	0.519	0.426	0.793	0.691	0.587	0.631	0.380	0.655	0.556	0.531	0.776	0.641
$\overline{M}PT^{GLS}$	14,568	10.955	60.181	28.140	26.899	10.337	44.49	30.039	31.174	26.600	10.595	30.067	28.899	18.822	44.194	28.727
Lag length	0	3	0	3	4	6	0	6	0	2	11	3	4	6	8	6
ADF	−0.470	−1.630	−1.280	−1.793	−0.334	−1.461	−1.23	−1.404	0.494	−0.723	−0.726	−0.584	0.181	−0.676	0.032	−0.470
Lag length	0	3	0	3	4	6	0	7	4	2	11	3	4	0	1	6
Range	1987:2–2008:4								1980:1–2008:4							
Critical values:																
Ng and Perron	$\overline{M}Z_{\alpha}^{GLS}$	$\overline{M}Z_{t}^{GLS}$	$\overline{M}SB^{GLS}$	$\overline{M}PT^{GLS}$					ADF	1987:2–2008:4			1980:1–2008:4			
1%*	−13.80	−2.580	0.174	1.780					1%*	−3.516			−3.490			
5%**	−8.100	−1.980	0.233	3.170					5%**	−2.899			−2.887			
10%***	−5.700	−1.620	0.275	4.450					10%***	−2.587			−2.581			

Notes: Test statistics defined in the text. "Lag length" refers to the lag length used in the $\overline{M}$ and ADF tests, respectively. The critical values are tabulated in Ng and Perron (2001). * Rejects null hypothesis at 1% significance level. ** Rejects null hypothesis at 5% significance level. *** Rejects null hypothesis at 10% significance level

Table 2 Unit root tests allowing for structural breaks

Dataset	Non-agricultural self-employment Seasonally adjusted data				Non-agricultural self-employment Non-seasonally adjusted data				Self-employment Seasonally adjusted data				Self-employment Non-seasonally adjusted data			
Country	US	Spain			US	Spain			US	Spain			US	Spain		
Status	S	S	E	OA	S	S	E	OA	S	S	E	OA	S	S	E	OA
(A)																
Min-*t*	−4.093	−3.750	−3.090	−4.133	−3.258	−3.542	−3.057	−4.040	−4.127	−3.482	−3.608	−3.261	−3.809	−3.593	−3.600	−3.444
	1997:2	1998:4	1991:1	1998:4	1997:2	1998:4	1999:3	1998:4	1997:2	1998:4	1991:1	1998:4	1997:3	1998:4	1991:1	1998:2
Lag length	0	3	0	3	4	4	1	4	0	3	0	3	4	4	1	4
(B)																
Min-*t*	−3.206	−2.316	−4.015	−2.402	−2.518	−2.856	−3.808	−2.545	−3.237	−2.346	−2.730	−2.977	−3.161	−2.713	−2.960	−3.048
	1999:3	1992:2	1996:2	2005:3	1991:2	1992:2	1996:2	2005:4	1993:2	2004:3	1996:4	2004:3	1993:3	1992:2	1989:4	2004:3
Lag length	0	3	0	3	4	4	1	4	0	3	0	3	4	4	1	4
(C)																
Min-*t*	−4.051	−3.332	−3.968	−3.971	−3.236	−3.131	−3.662	−3.850	−4.230	−2.860	−3.234	−3.201	−3.931	−3.094	−3.110	−3.496
	1997:2	1998:4	1999:3	1998:4	1997:2	1998:4	1995:2	1998:4	1997:2	1998:4	1991:1	1998:4	1997:2	2001:3	1989:4	1987:3
Lag length	0	3	0	3	4	4	1	4	0	3	0	3	4	4	1	4
Range	1987:2–2008:4								1980:1–2008:4							

Notes: Periods corresponding to min-t statistics are under each value. Critical values for the min-t are given by Zivot and Andrews (1992). Asterisks are as in Table 1. Min t-statistics are computed using sequential regressions over $1 <$ trend break $< T$ based on the following equations: (A): $\Delta x_t = \delta_0^A + \delta_1^A t + \delta_2^A DU + \alpha^A x_{t-1} + \sum_{j=1}^{k} \phi_j^A \Delta x_{t-j} + e_t$, (B): $\Delta x_t = \delta_0^B + \delta_1^B t + \delta_2^B DT + \alpha^B x_{t-1} + \sum_{j=1}^{k} \phi_j^B \Delta x_{t-j} + e_t$, (C): $\Delta x_t = \delta_0^C + \delta_1^C t + \delta_2^C DU + \delta_3^C DT + \alpha^C x_{t-1} + \sum_{j=1}^{k} \phi_j^C \Delta x_{t-j} + e_t$ where the dummy variables $\mathrm{DU}_t = 1$ and $\mathrm{DT}_t = t - \mathrm{TB}$ for $t > \mathrm{TB}$ and 0 otherwise, and TB denotes the period at which a possible trend break occurs. Critical values for the min-t are given by Zivot and Andrews (1992). In model (A): 1% (−5.34), 5% (−4.80), 10% (−4.58); Model (B): 1% (−4.93), 5% (−4.42), 10% (−4.11); Model (C): 1% (−5.57), 5% (−5.08), 10% (−4.82)

Table 3 Estimates of the linear unobserved component model

Dataset	Self-employment, agriculture excluded Seasonally adjusted data				Self-employment, agriculture excluded Non-seasonally adjusted data				Self-employment, agriculture included Seasonally adjusted data				Self-employment, agriculture included Non-seasonally adjusted data			
Country	US	Spain			US	Spain			US	Spain			US	Spain		
Status	S	S	E	OA	S	S	E	OA	S	S	E	OA	S	S	E	OA
Natural rate equation																
β	0.154	0.283**	0.404*	0.256**	0.400	0.297**	0.749*	0.218**	0.784*	0.175**	0.351**	0.205	0.613*	0.165**	0.320*	0.232*
	(0.115)	(0.163)	(0.115)	(0.113)	(0.693)	(0.170)	(0.049)	(0.093)	(0.037)	(0.087)	(0.149)	(0.129)	(0.127)	(0.081)	(0.071)	(0.053)
Cyclical rate equation																
φ_1	0.710*	1.462*	1.354*	1.529*	1.139	1.214*	1.264*	1.394*	1.177*	1.622*	1.063*	1.666*	1.139*	1.608*	1.063*	1.590*
	(0.194)	(0.213)	(0.131)	(0.133)	(0.784)	(0.182)	(0.048)	(0.117)	(0.018)	(0.111)	(0.060)	(0.146)	(0.013)	(0.114)	(0.092)	(0.043)
φ_2	−0.127***	−0.548*	−0.639*	−0.602*	−0.354	−0.378*	−0.523*	−0.500*	−0.658*	−0.671*	−0.326*	−0.711*	−0.644*	−0.680*	−0.298***	−0.645*
	(0.068)	(0.146)	(0.072)	(0.094)	(0.325)	(0.093)	(0.106)	(0.076)	(0.032)	(0.023)	(0.237)	(0.129)	(0.046)	(0.140)	(0.160)	(0.017)
σ_N	0.061*	0.150*	0.011*	0.147*	0.015*	0.022*	0.013*	0.020*	0.011	0.129*	0.010	0.122*	0.017*	0.119*	0.011	0.178*
	(0.021)	(0.014)	(0.001)	(0.013)	(0.001)	(0.002)	(0.001)	(0.002)	(0.007)	(0.016)	(0.007)	(0.014)	(0.001)	(0.026)	(0.077)	(0.016)
σ_C	0.086*	0.061*	0.125*	0.052*	0.002	0.013	0.000	0.027	0.001	0.082	0.088*	0.068*	0.000	0.093*	0.083**	0.080*
	(0.017)	(0.012)	(0.017)	(0.011)	(0.075)	(0.676)	(0.001)	(0.088)	(0.028)	(0.122)	(0.006)	(0.018)	(0.002)	(0.023)	(0.037)	(0.009)
f_o	0.026	0.025	0.088	0.027	0.046	0.025	0.080	0.027	0.121	0.023	0.069	0.025	0.124	0.036	0.037	0.022
$1/f_o$	38.18	39.66	11.3	36.63	21.59	39.36	12.38	37.12	8.28	43.48	14.47	40.34	8.04	28.04	27.31	44.63
Range	1987:2–2008:4								1980:1–2008:4							

Notes: Standard errors are in parentheses. Asterisks are as in Table 1

$f_o = (1/2\pi)\arccos\left[|\varphi_1|/2\left(-\varphi_2^{1/2}\right)\right]$

$1/f_o$ = Average peak-to-peak period, expressed in quarters

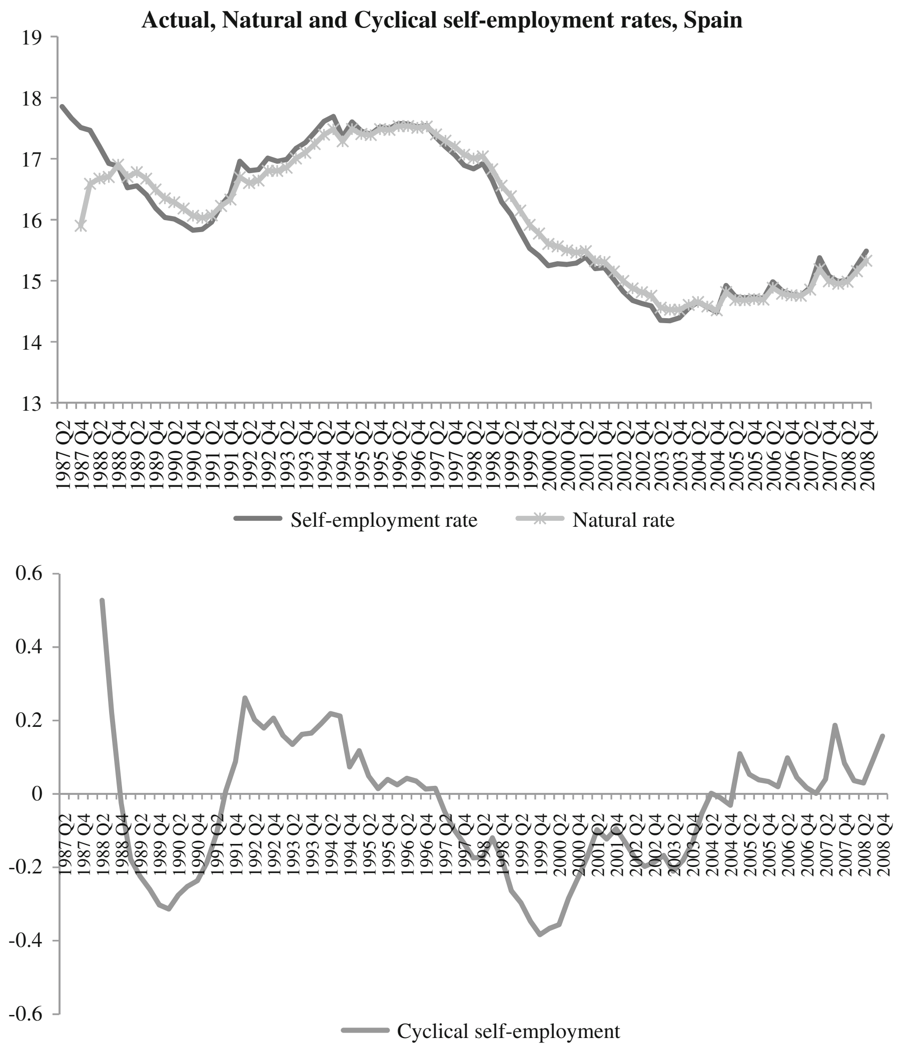

Fig. 1 Actual, natural, and cyclical self-employment rates, Spain

for the employer self-employment rate than for the own-account self-employment rate, suggesting that rates of employer self-employment are especially sensitive to shocks.

The second block of Table 3 (based on non-seasonally adjusted data) yields qualitatively similar results for the hysteresis parameter. The major differences come in the third and fourth blocks, which define self-employment to include the agricultural sector. Now the US as well as the Spanish estimates point to positive and statistically significant hysteresis. For Spain, hysteresis is again more pronounced amongst employers than amongst the own-account self-employed.

Finally, the sum of the estimates of the autoregressive parameters can be interpreted as another indication of persistence. It is noteworthy that the estimated parameters

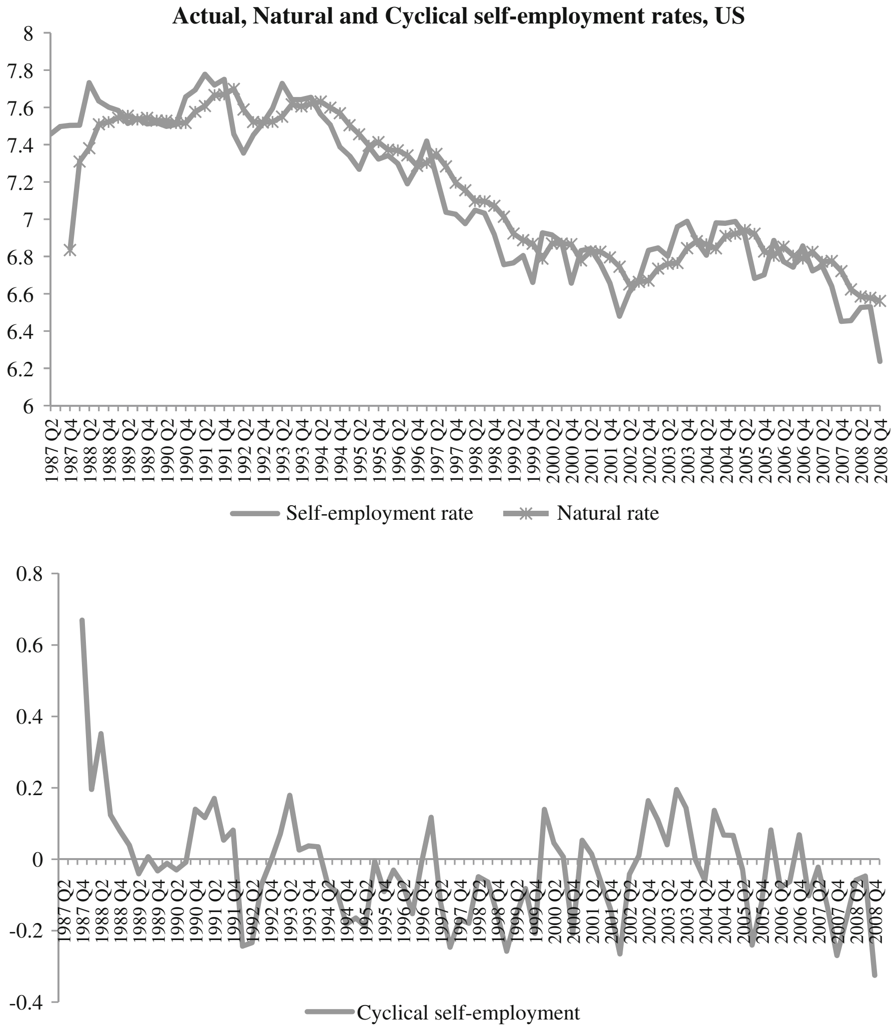

Fig. 2 Actual, natural, and cyclical self-employment rates, US

are consistent with average periodicities of the cyclical self-employment rate and own-account self-employment rate of about 9–10 years; and of about 3 years for the employer self-employment rate[15] (see the last two rows of Table 3).

[15] The spectral density of a stationary AR(2) process, when it has a peak, will have one at a frequency given by $f_{\mathrm{o}} = (1/2\pi)\arccos\left[|\varphi_1|/2\left(-\varphi_2^{1/2}\right)\right]$. As Jaeger and Parkinson demonstrated, the average peak-to-peak period is approximately $1/f_{\mathrm{o}}$. For example, using the results for the Spanish self-employment., ϕ_1= 1.462 and $\phi_2 = -0.548$, it follows that $f_{\mathrm{o}} = 0.025$ and $1/f_{\mathrm{o}}$ is about 38 quarters. Thus, the average periodicity of the cyclical component of Spanish self-employment is estimated to be just over 9 years.

Table 4 Testing for non-linearity

Country	Variable	Non-linear model Bootstrap p value
The US	Self-employment	0.310
Spain	Self-employment	0.750
	Employers	0.800
	Own-account workers	0.750

p values are obtained by using the method developed by Pérez-Alonso and Di Sanzo (2011)

4.3 Asymmetries

We next check whether our results are robust to the linear specification of the unobserved component model[16]. This involves jointly estimating the structure (4), (6), and (7) to determine whether there is a threshold for income growth which is associated with asymmetric business cycle responses. We wish to check whether the findings in the previous subsection are robust to possible asymmetries, or whether they were merely an artifact of the restrictive technical assumption of linearity.

Table 4 reports the p values calculated following the bootstrap technique described in detail in Pérez-Alonso and Di Sanzo (2011). The null hypothesis $H_0 : \beta^+ = \beta^-$ is not rejected for the aggregate self-employment series. Hence the hysteresis results for Spain in the linear model appear to be fairly robust in this regard.

4.4 Entrepreneurship and the business cycle

As discussed earlier, the correlation coefficient between the cyclical component of GDP and the self-employment rate can be used to infer pro- or counter-cyclical variations in self-employment, depending on whether this coefficient is positive or negative.

The estimates reported in the last row of Table 5 suggest that business cycle variations in output are significantly associated with cyclical aggregate self-employment rates in Spain—but not the USA. Separating out the aggregate self-employment series into its two components of employer and own-account self-employment in Spain generates an interesting additional finding which would otherwise be disguised: this correlation is positive for employers and negative for own-account self-employees. Thus, the Spanish employer self-employment rate appears to be pro-cyclical, while the own-account self-employment rate is anti-cyclical. These findings are consistent with our earlier conjecture that the most promising own-account self-employees switch to employer status in good times, while the least promising own-account workers are pulled into paid-employment as the demand for labor expands and employee wages rise.

[16] In this case, the assumption of stationarity should be tested by using an alternative method. We employed the Caner and Hansen (2001) methodology to test for a unit root in a TAR model. The null hypothesis is that there is not a unit root. The p values—calculated by using the bootstrap technique described in Pérez-Alonso and Di Sanzo (2011)—rejected the null for all time-series considered in this model.

Table 5 Estimates of the bivariate unobserved component model

Dataset Characteristics	Self-employment, agriculture excluded Seasonally adjusted data			
Country	US	Spain		
Parameter/bivariate model	GDP-S	GDP-S	GDP-E	GDP-OA
μ	0.698*	0.671**	0.505***	0.584*
	(0.077)	(0.161)	(0.258)	(0.142)
φ_{1y}	0.950*	1.224*	1.122*	0.734*
	(0.133)	(0.066)	(0.053)	(0.062)
φ_{2y}	−0.205	−0.396*	−0.216*	−0.526*
	(0.132)	(0.106)	(0.077)	(0.135)
φ_{1S}	0.346***	1.246*	1.251*	1.083*
	(0.201)	(0.182)	(0.169)	(0.021)
φ_{2S}	−0.159**	−0.429***	−0.269	−0.733*
	(0.070)	(0.258)	(0.180)	(0.113)
Cross series correlations	GDP-S	GDP-S	GDP-E	GDP-OA
Natural GDP/natural self-employment	−0.042	−0.403**	0.656*	−0.338*
	(0.248)	(0.202)	(0.196)	(0.130)
Natural GDP/cyclical self-employment	0.333	0.357	−0.769*	0.216
	(0.248)	(0.299)	(0.171)	(0.148)
Cyclical GDP/natural self-employment	0.110	0.599*	−0.580*	0.461*
	(0.220)	(0.181)	(0.217)	(0.154)
Cyclical GDP/cyclical self-employment	−0.425	−0.529**	0.674*	−0.345*
	(0.264)	(0.265)	(0.201)	(0.166)

Standard errors are in parentheses. Asterisks are as in Table 1

As noted earlier, the correlation between the natural components of GDP and the self-employment rate sheds light on this relationship in the long-run. Interestingly, the fourth row from the bottom of Table 5 points to an identical pattern as observed for the error correlations (cf., Sinclair 2009, in the case of unemployment and GDP). In the case of Spain, this implies that the counter-cyclical nature of entrepreneurship observed over the business cycle persists in the long-run too.

5 Conclusions

This article estimated unobserved components models for the self-employment rate in two developed but otherwise rather different economies: Spain and the United States. Defining hysteresis in terms of the interdependent evolution of a non-stationary natural rate and a stationary cyclical component, thereby distinguishing hysteresis from natural rate shocks, the results provide robust evidence of hysteresis in Spain but not in the USA. This implies that economic and/or policy shocks in Spain have permanent effects on rates of entrepreneurship. In view of the economic importance of entrepreneurship in modern economies, these results suggest that policy-makers may need to take particular care when designing pro-entrepreneurship and macroeconomic stabilization policies—especially in Spain. In view of evidence that pro-entrepreneurship policies can have unintended negative as well as positive side-effects on entrepreneurial outcomes (Parker 2009, Chapter 15), the case for public interventions therefore needs to be very compelling, since they can have profound long-run effects. An essential tool policy-makers need to make informed judgments in this regard is detailed policy evaluations. Our results argue for the use of much longer time horizons in formal evaluation exercises than the few years which are commonly used to gauge entrepreneurship policy impacts.

Our results also shed new light on the important but somewhat neglected issue of business cycle effects on entrepreneurship. Although we found some evidence of counter-cyclicality of entrepreneurship rates, deeper analysis showed that one should distinguish between own-account and employer components of self-employment. Employers comprise the minority of the self-employed in most countries, including Spain; but they are usually associated with greater economic value-added (Cowling et al. 2004). We found that in Spain employer self-employment rates evolve pro-cyclically whereas own-account self-employment rates evolve counter-cyclically. Therefore, the "quality" if not the quantity of entrepreneurship in Spain appears to evolve in a pro-cyclical manner.

In contrast to Spain, our results based on US data point to weak or non-existent hysteresis and business cycle effects in entrepreneurship. This might reflect the different nature of institutional and economic conditions in the US compared with Spain. However, we cannot rule out the possibility that it might also simply reflect data limitations, including the possibility that self-employment is an unsatisfactory practical measure of entrepreneurship. Further research is needed to determine whether it is different national (e.g., institutional) and economic conditions, or merely different data definitions of self-employment, which explain the diverse findings. Future work could also fruitfully apply the methodology used in this article to a broader range of countries, and should seek to lengthen the length of the data series that are utilized. A "micro" look at the causal processes underlying business cycle and hysteresis dynamics would be another natural extension of this article, complementing the "macro" analysis performed here. In the context of policy analysis, the impact of particular regulations or macro policies on entrepreneurial entry and longevity could be explored over long time periods, ideally within a natural experiment framework. That might help unlock the deep causes of hysteresis and business cycle effects, which were detected in this article, but which deserve much more detailed micro-level analysis to bridge the micro–macro divide.

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