Threshold cointegration and nonlinear adjustment between goods and services inflation: the case of Spain*

Vicente Esteve†
Departamento de Economía Aplicada II,
Universidad de Valencia, 46022 Valencia, Spain
José Antonio Martínez-Serrano
Departamento de Economía Aplicada II,
Universidad de Valencia, 46022 Valencia, Spain
Rafael Llorca-Vivero
Departamento de Economía Aplicada II,
Universidad de Valencia, 46022 Valencia, Spain

January 2005

Abstract

In this paper, we model the long-run relationship between goods and services inflation for the Spanish economy during the period 1971:2 to 2004:1. Our empirical methodology makes use of recent developments on threshold cointegration that consider the possibility of a nonlinear relationship between the two inflation series. According to our results, the null hypothesis of linear cointegration would be rejected in favor of a two-regime threshold cointegration model. Consequently, we could expect a cointegrating relationship only when the divergence between services inflation and goods inflation is above the threshold point estimate.

Keywords: Inflation, goods, services, threshold cointegration.

JEL classification: C32, E31.

*The authors acknowledge financial support from the Spanish Ministry of Science and Technology, through the projects SEC2002-03651 (V. Esteve) and SEC2003-05836 (J.A. Martínez-Serrano and R. Llorca-Vivero).
† Corresponding author: Vicente Esteve, Departamento de Economía Aplicada II, Universidad de Valencia, 46022 Valencia (Spain). Tel.: +34-96-3828349; fax: +34-96-3828354. E-mail: vicente.esteve@uv.es.
1 Introduction

A noteworthy feature of the behavior of the inflation rate is the divergence between the rate of increase of services prices and that of goods prices. Prices of services tend to increase faster than goods prices. Consequently, the behavior of the services sector is very often considered as the main source of inflationary pressures. However, this will be a problem only if the gap between services and goods prices widens in the long-run.

This question has been addressed for the US economy by Peach, Rich and Antoniades (2004), who using quarterly data for the period 1967-2002 find a linear cointegrating relationship between both series of data. That is, the authors show that the gap between inflation rates has a strong tendency to a "constant equilibrium value" in the long-run. In the USA case, the equilibrium is restored through a rise in goods inflation and a decrease in services inflation.

In this paper, we investigate the long-run relationship between goods and services prices for the Spanish case. Our empirical methodology makes use of recent developments on threshold cointegration that consider the possibility of a nonlinear relationship between the two inflation series. In the empirical application we use quarterly data over the period 1971-2004. To preview the results, we find a two-regime threshold cointegration (non-linear cointegration) between goods and services inflation in which differences beyond 3.0 percentage points drive to a downward pressure in services inflation in order to restore the equilibrium.

The rest of this paper is organized as follows: Section 2 reports unit root tests; Section 3 presents the nonlinear cointegration testing strategy and reports threshold cointegration test results and Section 4 concludes.

2 Data and unit root tests

The top panel of Figure 1 plots Spanish goods inflation and services inflation for the period 1971:2-2004:1. For goods, we have taken yearly changes of the quarterly implicit GDP deflator of agriculture and industrial branches jointly. Therefore, energy and construction branches of activity are excluded. For services, we have performed the same calculations for market services. Data of both series come from Financial Accounts of the Spanish Economy (Bank of Spain, http://www.bde.es/estadis/ccffe/ccffe.htm).

Except for some specific years, services inflation is higher than goods inflation, with average values of 10.04% and 6.92% over the sample period, respectively. The average gap is, therefore, 3.12%. Although a similar tendency is observed in both series, it is easy to detect some periods of huge differentials (bottom panel). Additionally, we do not detect any cyclical pattern over the period, that is, there is no correspondence between periods of recession or expansion with periods in which the gap narrows or widens. Anyway, it is not possible to say that the gap behaves in a steady manner, which suggests that the series do not probably exhibit a linear cointegrating relationship. We will test the diverse possibilities in the next sections. This approach will allow us to consider the possibility of a nonlinear long-run relationship between goods and services inflation, so that a mean-reverting dynamic behavior of the "gap" between the two inflation rates (or a long-run equilibrium relationship) should
be expected only once a certain threshold is reached.

As a first step of the analysis, we have tested for the order of integration of the two series. Ng and Perron (1995) have recently developed a set of state-of-the-art unit root tests that have good size and power properties. This methodology consists of a class of modified tests, \( M_{Z_G}^{GLS} \), \( M_{S_G}^{GLS} \) and \( M_{Z_i}^{GLS} \), 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 criterion \( (MAIC) \). In addition, Ng and Perron (2001) have proposed a similar procedure that corrects the problems associated with the standard Augmented Dickey-Fuller test, \( ADF^{GLS} \). Table 1 displays the results of Ng and Perron tests. As shown, the null hypothesis of non stationarity for the two series in levels cannot be rejected, independently of the test. Accordingly, the two series would be concluded to be I(1).

3 Threshold cointegration

3.1 Methodology

The concept of threshold cointegration was introduced by Balke and Fomby (1997) as a feasible way to combine nonlinearity and cointegration. Balke and Fomby (1997) stressed the possibility that this movement towards the long-run equilibrium might not occur in every time period, due to the presence of some adjustment costs on the side of economic agents. In other words, there could be a discontinuous adjustment to equilibrium so that, only when the deviation from the equilibrium exceeds a critical threshold, the benefits of adjustment are higher than the costs, and economic agents move the system back to equilibrium. Threshold cointegration would characterize this discrete adjustment as follows: the cointegrating relationship does not hold inside a certain range, but holds if the system gets ‘too far’ from the equilibrium; i.e., cointegration would hold only if the system exceeds a certain threshold.

Recently, Hansen and Seo (2002) have contributed further to this literature by examining the case of an unknown cointegration vector. In particular, these authors proposed a vector error-correction model (VECM) with one cointegrating vector and a threshold effect based on the error-correction term, and developed a Lagrange multiplier (LM) test for the presence of a threshold effect. Hansen and Seo (2002) considered a two-regime threshold cointegration model, or a nonlinear VECM of order \( l + 1 \), such as:

\[
\Delta x_t = \begin{cases} 
A_1' X_{t-1}(\beta) + u_t & \text{if } w_{t-1}(\beta) \leq \gamma \\
A_2' X_{t-1}(\beta) + u_t & \text{if } w_{t-1}(\beta) > \gamma
\end{cases} \tag{1}
\]

where \( x_t \) is a \( p \)-dimensional \( I(1) \) time series which is cointegrated with one \( p \times 1 \) cointegrating vector \( \beta \), \( w_t(\beta) = \beta' x_t \) is the \( I(0) \) error-correction term, \( u_t \) is an error term, \( A_1 \) and \( A_2 \) are coefficient matrices, and \( \gamma \) is the threshold parameter. As can be seen, the threshold model (1) has two regimes, depending on whether deviations from the equilibrium (defined by the value of the error-correction term) are below or above the threshold, where \( A_1 \) and \( A_2 \) describe the dynamics in each of the regimes. In one of the regimes there would be no tendency for the variables \( x_t \) to revert to an equilibrium (i.e., the variables would not be cointegrated); on the contrary, in the other regime there would
be a tendency for the variables $x_t$ to move towards some equilibrium (i.e., the variables would be cointegrated).

### 3.2 Results

First, we tested for the presence of linear cointegration using two of the most common and straightforward linear cointegration tests: the OLS-residual based test of Engle and Granger (1987), and Phillips and Ouliaris (1990). The first step in each test involves estimating the potential cointegration relationship between services inflation, $\pi^\text{Services}_t$, and goods inflation, $\pi^\text{Goods}_t$, with cointegrating vector $(1, -\beta)'$. The second step of each test is to examine the stationarity properties of the OLS residuals. Engle and Granger (1987) test the stationarity properties of the cointegrating residuals using the ADF test, whereas Phillips and Ouliaris (1990) use the $Z_a$ test. The null hypothesis is no linear cointegration and the alternative hypothesis is linear cointegration. To apply the $Z_a$ test we use the semiparametric fully modified estimator of Phillips Hansen (1990) (PHFM). To estimate the linear cointegrating relationship we also use the Dynamic Ordinary Least Squares (DOLS) estimation of Stock and Watson (1993) within the methodology proposed by Shin (1994). The null hypothesis is linear cointegration and the alternative is no linear cointegration. The first step consists of the estimation of a long-run dynamic equation including leads and lags of the explanatory variables. The second step of the Shin test is based on the calculation of a Lagrange statistic ($LM$), $C_p$.

In Table 2 we present the results of the linear cointegration tests. According to the result of the ADF test in the long-run relationship, it is not possible to reject the null of no cointegration. For the case of the $Z_a$ test there is evidence of linear cointegration between $\pi^\text{Services}_t$ and $\pi^\text{Goods}_t$ but only at the 10% significance level. However, according to the result of the $C_p$ test the null of linear cointegration is not rejected at the 1% level. From the various tests applied, we can conclude that there is not clear evidence of linear cointegration between goods and services inflation. In all cases, the estimated cointegrating parameter is quite close to a unit coefficient ($\beta = 1$). The reasons that explain these disappointing results may be the presence of non-linearities in inflation rates series. Some authors have found that inflation is a nonlinear regime-dependent process, see Kim (1993), Evans and Lewis (1995), and Henry and Shields (2004), inter alia.

Next, we explore the possibility that a threshold cointegration provides a better empirical description of the long-run relationship between goods and services inflation. To address this question, we may write the linear regression model as a bivariate linear cointegrating VAR model with one lag, $l = 1$, such as:

$$
\begin{pmatrix}
\Delta \pi^\text{Services}_t \\
\Delta \pi^\text{Goods}_t
\end{pmatrix} = \mu + \delta w_{t-1} + \Gamma \begin{pmatrix}
\Delta \pi^\text{Services}_{t-1} \\
\Delta \pi^\text{Goods}_{t-1}
\end{pmatrix} + \varepsilon_t
$$

(2)

where the long-run relationship is defined as $w_{t-1} = \pi^\text{Services}_{t-1} - \beta \pi^\text{Goods}_{t-1}$ with cointegrating vector $(1, -\beta)'$.

As the parameter estimates were quite close to a unit coefficient (see Table 2), we have applied the test of threshold cointegration sup $LM^\beta$ (for a given $\beta = 1$) proposed by Hansen and Seo (2002). The $p$-values are calculated using a
and services inflation. According to our results, the null hypothesis of linear integration that consider the possibility of a nonlinear relationship between goods and services inflation would be strongly rejected. The estimated threshold is $\hat{\gamma} = 3.0$, with the error-correction term defined as $w_t = \pi_t^{\text{Services}} - \pi_t^{\text{Goods}}$ (i.e., the gap between the two inflation rates). Hence, the first regime (including 45% of the observations) would occur when the services inflation are less than 3.0 percentage points above the goods inflation; in other words, when the gap is below 3.0%. In turn, the second regime (with 55% of the observations) would occur when the gap is above 3.0%.

On the other hand, the corresponding two-regime threshold VAR (with heteroskedasticity-consistent standard errors in parentheses) is:

$$\Delta \pi_t^{\text{Services}} = \begin{cases} -0.002 - 0.04 \ w_{t-1} + 0.69 \ \Delta \pi_{t-1}^{\text{Services}} + 0.02 \ \Delta \pi_{t-1}^{\text{Goods}} + u_{1t}, \ w_{t-1} \leq 3.0, \\ 0.05 - 0.16 \ w_{t-1} + 0.36 \ \Delta \pi_{t-1}^{\text{Services}} + 0.06 \ \Delta \pi_{t-1}^{\text{Goods}} + u_{2t}, \ w_{t-1} > 3.0, \end{cases}$$

$$\Delta \pi_t^{\text{Goods}} = \begin{cases} -0.67 + 0.30 \ w_{t-1} + 0.21 \ \Delta \pi_{t-1}^{\text{Services}} + 0.43 \ \Delta \pi_{t-1}^{\text{Goods}} + u_{1t}, \ w_{t-1} \leq 3.0, \\ 0.50 - 0.08 \ w_{t-1} + 0.02 \ \Delta \pi_{t-1}^{\text{Services}} - 0.04 \ \Delta \pi_{t-1}^{\text{Goods}} + u_{2t}, \ w_{t-1} > 3.0, \end{cases}$$

The estimation of the error-correction term in the VAR, $w_{t-1}$, allows for a straightforward investigation into the behavior of the gap between services and goods inflation in the Spanish economy. We can also examine the sign and magnitude of these coefficients in order to analyze the adjustment process by which long-run equilibrium between the inflation series is restored. First, there is a strong and statistically significant error-correction term in the services inflation equation when $\pi_t^{\text{Services}}$ is above $\pi_t^{\text{Goods}}$ (or the gap is relatively high) and in the goods inflation equation when $\pi_t^{\text{Services}}$ is below $\pi_t^{\text{Goods}}$ (or the gap is relatively low). Second, a value of the gap above 3.0% in one quarter will produce downward pressure on services inflation in the subsequent quarter. Third, a value of the gap below 3.0% in one quarter will produce upward pressure on goods inflation in the subsequent quarter. The coefficients on the error-correction term also indicate that the magnitude of the response for goods inflation is twice that of the services inflation.

4 Conclusions

In this paper, we model the long-run relationship between goods and services inflation for Spanish economy during the period from 1971:2 to 2004:1. Our empirical methodology makes use of recent developments on threshold cointegration that consider the possibility of a nonlinear relationship between goods and services inflation. According to our results, the null hypothesis of linear cointegration would be rejected in favor of a two-regime threshold cointegration.
model. Consequently, a system of two regimes would seem to characterize the discontinuous or nonlinear adjustment of the services inflation rates towards a long-run equilibrium, with the threshold parameter estimated at 3.0 percentage points. Therefore, we could expect a cointegrating relationship only when the divergence between services inflation and the adjustment for goods inflation is above 3.0%.

Furthermore, when the gap is above the estimated threshold parameter, a decrease in services inflation ought to occur in order to restore the long-run equilibrium between the two inflation series. In contrast, when the gap is below the estimated threshold parameter, a rise in services inflation would have to occur in order to restore the long-run equilibrium between the two inflation series.

References


Table 1
Ng-Perron tests for a unit roots\textsuperscript{a,b}

I(1) vs. I(0) Case: $p = 1, \hat{c} = -13.5$

<table>
<thead>
<tr>
<th>Variable</th>
<th>$M Z_i^{GLS}$</th>
<th>$M Z_i^{GLS}$</th>
<th>$M S B^{GLS}$</th>
<th>$A D F^{GLS}$</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\pi_t^{Goods}$</td>
<td>-4.31</td>
<td>-1.46</td>
<td>0.340</td>
<td>-1.38</td>
</tr>
<tr>
<td>$\pi_t^{Services}$</td>
<td>-6.15</td>
<td>-1.74</td>
<td>0.284</td>
<td>-1.54</td>
</tr>
</tbody>
</table>

Notes:
\textsuperscript{a} A *, ** and *** denote significance at the 10%, 5% and 1% levels, respectively.

\textsuperscript{b} The MAIC information criteria is used to select the autoregressive truncation lag, $k$, as proposed in Perron and Ng (1996). The critical values are taken from Ng and Perron (2001), table 1.
<table>
<thead>
<tr>
<th>Methodology</th>
<th>$\beta$</th>
<th>Test</th>
<th>Method</th>
</tr>
</thead>
<tbody>
<tr>
<td>Engle and Granger (1987)$^a$</td>
<td>0.97</td>
<td>$-2.62$</td>
<td>OLS</td>
</tr>
<tr>
<td>Phillips and Ouliaris (1990)$^b$</td>
<td>1.05</td>
<td>$-19.55^*$</td>
<td>PHPM</td>
</tr>
<tr>
<td></td>
<td>0.97</td>
<td>$-18.13^*$</td>
<td>OLS</td>
</tr>
<tr>
<td>Stock and Watson (1993)$^c$</td>
<td>1.13</td>
<td>0.074</td>
<td>DOLS</td>
</tr>
</tbody>
</table>

Notes:

$^a$ We follow Ng and Perron (1995) and select the lag order $k$ via top-down testing, considering a maximum lag order of 12. The estimated model is:

$$\pi_t^{Services} = \alpha + \beta \pi_t^{Goods} + \eta_t^d.$$

$^b$ The number of lags truncation parameter selected for the Barlett window has been $l = 11 \approx \int \left[ T^{1/2} \right]$, as proposed in Newey and West (1987).

$^c$ The number of leads and lags selected has been $q = 5 \approx \int \left[ T^{1/3} \right]$, as proposed in Stock and Watson (1993). The long-run variance of the cointegrating regression residual has been estimated using the Bartlett window with $l = 11 \approx \int \left[ T^{1/2} \right]$, as proposed in Newey and West (1987). The estimated model is:

$$\pi_t^{Services} = \alpha + \beta \pi_t^{Goods} + \sum_{j=q}^{q} \varphi_j \Delta \pi_{t-j}^{Goods} + \nu_t.$$

$^d$ A *, ** and *** denote significance at the 10%, 5% and 1% levels, respectively. The critical values are taken from MacKinnon (1991) ($ADF$), table 1, $N = 2$; from Haug (1992) ($Z_o$), Table 2, $m = 1$, $T = 150$; and from Shin (1994) ($C_\mu$).

<table>
<thead>
<tr>
<th>Critical values:</th>
<th>10%</th>
<th>5%</th>
<th>1%</th>
</tr>
</thead>
<tbody>
<tr>
<td>$ADF$</td>
<td>-3.07</td>
<td>-3.38</td>
<td>-3.98</td>
</tr>
<tr>
<td>$Z_o$</td>
<td>-16.54</td>
<td>-19.83</td>
<td>-27.40</td>
</tr>
<tr>
<td>$C_\mu$</td>
<td>0.231</td>
<td>0.314</td>
<td>0.533</td>
</tr>
</tbody>
</table>
Table 3
Hansen-Seo tests of threshold cointegration

<table>
<thead>
<tr>
<th></th>
<th>( l = 1 )</th>
<th>( l = 2 )</th>
</tr>
</thead>
<tbody>
<tr>
<td>Test statistic value</td>
<td>22.09</td>
<td>24.59</td>
</tr>
<tr>
<td>Calculated ( p )-values</td>
<td>0.016</td>
<td>0.060</td>
</tr>
<tr>
<td>Threshold parameter</td>
<td>3.01</td>
<td>2.94</td>
</tr>
<tr>
<td>Estimate of the cointegrating vector</td>
<td>1.00</td>
<td>1.00</td>
</tr>
</tbody>
</table>

Note:
The \( \text{sup}\, LM^0 \) is a heteroskedastic-consistent LM test statistic for the null hypothesis of linear cointegration (i.e., there is no threshold effect), against the alternative of threshold cointegration (i.e., model (1)), where \( \bar{\beta}_0 \) is the known value of \( \beta \) (in the case analyzed below, \( \bar{\beta}_0 = 1 \)), and \([\gamma_L, \gamma_U]\) is the search region set so that \( \gamma_L \) is the \( \pi_0 \) percentile of \( \hat{\epsilon}_{t-1} \), and \( \gamma_U \) is the \( (1 - \pi_0) \) percentile.