Why should Central Banks avoid the use of the underlying inflation indicator?

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The analyses, opinions and findings of this paper represent the views of the authors, they are not necessarily those of the Banco de Portugal.

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Abstract

This paper assesses the usefulness of the commonly used underlying inflation indicator, in light of the criteria proposed in Marques et al. (2000). Empirical evidence for a group of six countries strongly suggests that the use of underlying inflation as an indicator of trend inflation should be avoided.

Keywords: core inflation indicator; underlying inflation; evaluation criteria

JEL classification: C43, E31, E52

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1. Introduction

The “CPI excluding unprocessed food and energy” indicator, also known as “underlying inflation”, was one of the first core inflation indicators ever proposed in the literature. Blinder (1982) uses the CPI excluding food, energy and mortgage interest costs to estimate the underlying inflation for the USA in the 1970s, as these components were largely responsible for the inflationary shocks of 74 and 78-80. During the 1980s and early 1990s, this type of indicators became extremely popular amongst Central Banks.¹ The procedure is motivated by the high volatility of the excluded categories, which is supposed to be caused by temporary, non-monetary phenomena.²

However, recent research on core inflation has cast some doubts on the usefulness of this indicator. Other core inflation indicators have been developed – such as trimmed means – and most of them seem to outperform the underlying inflation measure. See, for instance, Bryan and Cecchetti (1994), Freeman (1998) and Marques et al. (2000). On the other hand, since there is no economic model explaining the construction of the indicator, the selection of the excluded items is a purely subjective decision. From a theoretical point of view, although we can think of core inflation estimation as a signal extraction problem, it seems highly unrealistic to assume that some categories contain no information at all.

This paper presents further evidence against the use of the underlying inflation indicator, resorting to the concept of desirable properties of a measure of core inflation, as defined in Marques et al. (2000). The empirical evidence presented below shows that the underlying inflation indicator systematically fails these conditions, for a reasonably large number of different countries.

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¹ The monitoring of such an indicator is common practice in the United States. Banco de Portugal [see Nascimento (1990)], Banco de España, Banca d’Italia, the Deutsche Bundesbank, De Nederlandsche Bank and Banque Nationale de Belgique have, among others, commented either on a regular basis or in a case-by-case basis the evolution of such a type of indicator. See Álvarez and Matea (1999) for a complete list of references. Finally, the European Central Bank regularly mentions a similar measure of inflation in its Monthly Bulletins.

² Empirical evidence that unprocessed food and energy prices are more volatile, on average, than the remaining components of the CPI is provided for instance in Álvarez and Matea (1999). Bryan et al. (1997) also provides empirical evidence that fresh food and energy goods are more likely to be trimmed than the average component of the CPI in the computation of their selected trimmed inflation measure.
This paper is organised as follows. Section 2 describes some conditions that a core inflation indicator should verify. In section 3, some arguments against the use of the underlying inflation indicator are presented. The empirical results are shown in section 4. Section 5 concludes.

2. The properties of core inflation indicators

Assume that for any given period \( t \) the inflation rate, say \( \pi_t \), is broken down into the sum of two components: a permanent component named core or trend inflation, say \( \pi_t^c \), and a temporary component represented by \( u_t \). Therefore, we have:

\[
\pi_t = \pi_t^c + u_t. \tag{1}
\]

In equation (1), we assume that the temporary disturbances in the inflation rate, \( u_t \), are caused by developments such as changes in weather conditions, disturbances in the demand and supply of goods, etc. By definition, \( u_t \) is expected to have zero mean and finite variance, and therefore non-stationarity is excluded on theoretical grounds. Notice for instance that, if \( u_t \) were allowed to exhibit a nonzero mean, then \( \pi_t^c \) would not be capturing the whole systematic component of \( \pi_t \).

According to Marques et al. (2000), when inflation is an I(1) process, an appropriate measure of core inflation, say \( \pi_t^* \), should verify the following:

i) \( \pi_t^* \) is I(1) and \( \pi_t \) and \( \pi_t^* \) are cointegrated with unitary coefficient, i.e. \( (\pi_t - \pi_t^*) \) is stationary with zero mean;

ii) there is an error correction mechanism for \( \pi_t \) given by \( z_{t-1} = (\pi_{t-1} - \pi_{t-1}^*) \), i.e. \( \gamma \neq 0 \) in

\[z_{t-1} = \gamma (\pi_{t-1} - \pi_{t-1}^*) + \varepsilon_{t-1} \]

3 Marques et al. (2000) also proposes a set of testable conditions when the inflation rate is a stationary variable.

4 This condition was first proposed by Freeman (1998).
\[\Delta \pi_i = \sum_{j=1}^{m} \alpha_j \Delta \pi_{i-j} + \sum_{j=1}^{m} \beta_j \Delta \pi_{i-j}^* - \gamma (\pi_{i-1} - \pi_{i-1}^*) + \varepsilon_i; \quad (2)\]

iii) \(\pi_i^*\) is strongly exogenous for the parameters of equation (2).

Condition i) is a direct consequence of the definition of \(u_t\) in equation (1). Condition ii) may be interpreted as the requirement of \(\pi_i^*\) being an attractor for \(\pi_i\), as the error correction mechanism forces inflation to converge towards its trend. In other words, \(\pi_i^*\) should act as a leading indicator for \(\pi_i\). Condition iii) guarantees that the path of \(\pi_i^*\) is not influenced by past values of \(\pi_i\), i.e. that \(\pi_i\) does not Granger cause \(\pi_i^*\).

We test condition i) in two steps. First, we use a unit root test to establish the stationarity of \(z_t = (\pi_t - \pi_t^*)\). We then test the null \(\alpha = 0\) in the static regression

\[\pi_t = \alpha + \pi_t^* + u_t, \quad (3)\]
given that \(z_t\) is stationary. After establishing condition i), the verification of ii) is simple, just requiring the estimation of model (2). The hypothesis \(\gamma = 0\) can then be tested with the conventional t-ratio of \(\hat{\gamma}\). Condition iii) implies that in the error correction model for \(\pi_i^*\),

\[\Delta \pi_i^* = \sum_{j=1}^{r} \delta_j \Delta \pi_{i-j}^* + \sum_{j=1}^{r} \theta_j \Delta \pi_{i-j} - \lambda (\pi_{i-1} - \pi_{i-1}^*) + \eta_i, \quad (4)\]

the null hypothesis \(\lambda = \theta_1 = \ldots = \theta_r = 0\) should not be rejected. A necessary condition for iii) is weak exogeneity of \(\pi_i^*\). This can be verified by testing \(\lambda = 0\) in equation (4).
3. Why should one not expect underlying inflation to measure core inflation

The underlying inflation indicator is obtained by excluding some components of the CPI. Let us define

\[ P_t = \alpha P^0_t + (1 - \alpha) P^i_t \]  

where \( P_t \) stands for the CPI, \( P^i_t \) for the items excluded from the CPI with the argument that they are more volatile, i.e. energetic products and unprocessed food, and \((1 - \alpha)\) for the weight of the remaining goods and services. Therefore, by definition, \( P^0_t \) represents the price index used to compute underlying inflation. For monthly data this equation may be rewritten as

\[ \pi_t = \alpha w_t + (1 - \alpha)v_t \]  

where

\[ \pi_t = \frac{P_t}{P_{t-12}} - 1; \quad w_t = \frac{P^0_t}{P^0_{t-12}} - 1; \quad v_t = \frac{P^i_t}{P^i_{t-12}} - 1; \quad \alpha_t = \alpha \cdot \frac{P^0_t}{P^0_{t-12}}. \]  

Notice that \( w_t \) is our conventional underlying inflation indicator. Now we may ask under what conditions does \( w_t \) meet the first criterion for a core inflation indicator. Recall that this will be satisfied if, in the static regression

\[ \pi_t = \beta w_t + u_t \]  

we have \( \beta = 1 \) and \( u_t \sim I(0) \). Let us see under what circumstances we can approximate (6) by (8). If we define

\[ \alpha_t = \mu + \varepsilon_t \]
expression (6) can be rewritten as

$$\pi_t = \mu w_t + (1-\mu)v_t + (\alpha_t - \mu)(w_t - v_t) \; ,$$

(10)

Now, if \( \pi_t \) is I(1), one expects \( w_t \) and \( v_t \) also to be I(1).\(^5\) If \( v_t \) and \( w_t \) are not cointegrated, then clearly \( \pi_t \) and \( w_t \) may not be cointegrated either. On the other hand, if \( v_t \) and \( w_t \) are cointegrated, we may write

$$v_t = \theta w_t + \eta_t \; ,$$

(11)

with \( \eta_t \sim I(0) \). Inserting (11) into (10), one gets

$$\pi_t = [\mu + (1-\mu)\theta]w_t + u_t \; ,$$

(12)

where

$$u_t = (\alpha_t - \mu)(w_t - v_t) + (1-\mu)\eta_t \; .$$

(13)

Now, if \( \theta = 1 \), we are back to our definition of a core inflation measure presented in equation (1) of the previous section. So, in order to have \( \beta = 1 \) and \( u_t \sim I(0) \) in (8), it is necessary and sufficient that both \( \alpha_t \) and \( (w_t - v_t) \) are stationary variables.

The fact that this may not occur means that the items excluded from the CPI in order to compute \( w_t \), i.e. \( P_t^1 \), may contain some information that systematically differs from the one included in \( z_t \). Therefore, in computing the underlying inflation indicator, we may be excluding too much information, i.e., we may be excluding from \( \pi_t \) not only “noise” but also “signal”. If this is the case, \( w_t \) will not meet condition i) in section 2.

It is also easy to understand why \( w_t \) must not be expected to meet conditions ii) and iii), i.e. to be a leading indicator for \( \pi_t \) and not to be Granger caused by \( \pi_t \). In order to compute \( w_t \), we exclude from the CPI the prices of goods that enter as

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\(^5\) If \( v_t \sim I(0) \), then (6) states that \( \pi_t - \alpha_j w_t \) is stationary. Therefore, it is not possible to have \( \beta = 1 \) in (8), as \( 0 < \alpha_j < 1 \).
intermediate inputs in the production process (energetic goods and unprocessed food). Therefore, changes in $v_t$ are expected to direct and contemporaneously affect $\pi_t$ while affecting $w_t$ indirectly with a lag. This being so, $v_t$ is a leading indicator for $w_t$, and as long as it affects $\pi_t$, this means that $\pi_t$ also appears as a leading indicator for $w_t$. In practice, this situation would cause conditions ii) and iii) of the previous section not to be verified, since CPI inflation would cause, rather than be caused by, underlying inflation. As section 4 shows, this is the type of results obtained for the six countries considered in this study.

4. Empirical results

This paper assesses the empirical properties of underlying inflation for six different countries: USA, Germany, France, Italy, Spain and Portugal. For the cases of the USA and France, however, there is no available information for the unprocessed component of food and, therefore, the whole class of food was excluded. Data is described in Table 1.

<table>
<thead>
<tr>
<th>Series</th>
<th>Excluded items</th>
<th>Sample period</th>
</tr>
</thead>
<tbody>
<tr>
<td>USA</td>
<td>Food and energy</td>
<td>1987:1 2000:2</td>
</tr>
<tr>
<td>Germany</td>
<td>Seasonal food and energy</td>
<td>1992:1 2000:4</td>
</tr>
<tr>
<td>France</td>
<td>Food, energy and public utilities</td>
<td>1987:1 1997:12</td>
</tr>
<tr>
<td>Italy</td>
<td>Fresh food and energy</td>
<td>1987:1 2000:5</td>
</tr>
<tr>
<td>Spain</td>
<td>Unprocessed food and energy</td>
<td>1987:1 2000:2</td>
</tr>
<tr>
<td>Portugal</td>
<td>Unprocessed food and energy</td>
<td>1987:1 1999:12</td>
</tr>
</tbody>
</table>

Preliminary testing showed that the null of a unit root in CPI inflation could not be rejected (by an ADF test) for any country. Therefore, we are able to use the testing
procedure described in section 2. Table 2 contains the main results of these tests. The
stationarity tests on CPI inflation are shown in column 1. Columns 2 and 3 report the
tests on condition i). Column 4 refers to condition ii), while columns 5 and 6 present the
results on both versions of condition iii). Finally, column 7 shows the conclusions
drawn from the tests.

Since the tests for conditions ii) and iii) are conditional on the verification of i),
we did not test conditions ii) and iii) for the series that failed the first criterion (i.e. Italy
and Portugal). For Portugal, we conclude that the series are not cointegrated with
unitary coefficient, so it may be possible for CPI inflation to diverge from underlying
inflation for substantial periods of time. For Italy, in spite of evidence showing a strong
cointegration relationship, the test on column 3 reports a systematic bias of the
underlying inflation measure, which naturally lessens its interest as a core inflation
indicator.

For all other countries, condition i) is verified; however, condition ii) does not
hold. Recall that this condition required the underlying inflation indicator to “attract”
CPI inflation. The fact that this does not occur means that knowing CPI inflation is, say,
below underlying inflation in a given period, does not convey any information on the
future path of CPI inflation.

Given the results so far, one would of course expect condition iii) not to hold also.
Since \( \pi_t \) and \( \pi_t^* \) are cointegrated and there is no ECM representation for \( \pi_t \), the
Granger Representation Theorem requires that an ECM representation exists for \( \pi_t^* \).
This is in fact verified (in column 5) for all countries but Spain. However, even in this
case, the null would not be rejected at the 10% level of significance. This means that
underlying inflation is not a leading indicator for inflation and, moreover, that it is the
inflation rate itself that appears to lead the so-called indicator of underlying inflation, as
the analysis in section 3 suggests.

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6 In the testing procedure, we used significance levels of 10% for the ADF tests, and 5% for the \( t \) and \( F \) tests. We set the
orders of the lag polynomials in the ADF and the ECM regressions such that the residuals were not autocorrelated. Although the
ECM models presented in section 2 (equations (2) and (4)) do not include a constant term, we also tested conditions ii) and iii) with
a nonzero constant. We show the results for this case only when the conclusions differ from the main test.
5. Conclusions

This paper shows that the so-called measure of underlying inflation, which is used by several Central Banks as a measure of trend inflation, does not meet the necessary conditions set out in Marques et al. (2000). These conditions posit that any core inflation measure should be cointegrated with inflation (with a unit coefficient) and act as an attractor for inflation, i.e., to Granger cause inflation but not to be Granger caused by it. Therefore, it appears to be inappropriate to use this indicator to analyse the current status of inflation or to make inference about its likely future path.

References


Table 2 – Evaluating the underlying inflation indicator (a)

<table>
<thead>
<tr>
<th>Country</th>
<th>Stationarity of $\pi$</th>
<th>Stationarity of $(\pi - \pi^*)$</th>
<th>$\alpha = 0$ given $\beta = 1$</th>
<th>$\gamma = 0$</th>
<th>$\lambda = 0$</th>
<th>Strong exogeneity $\lambda = \theta_1 = \ldots = \theta_r = 0$</th>
<th>Conclusion</th>
</tr>
</thead>
<tbody>
<tr>
<td>USA</td>
<td>No</td>
<td>Yes</td>
<td>Yes</td>
<td>Yes</td>
<td>No (b)</td>
<td>No (c)</td>
<td>Fails conditions ii) and iii)</td>
</tr>
<tr>
<td></td>
<td>ADF(1) = -2.00</td>
<td>ADF(1) = -2.81 *</td>
<td>P = 0.163</td>
<td>P = 0.250</td>
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<tr>
<td>Germany</td>
<td>No</td>
<td>Yes</td>
<td>Yes</td>
<td>Yes</td>
<td>No</td>
<td>No</td>
<td>Fails conditions ii) and iii)</td>
</tr>
<tr>
<td></td>
<td>ADF(1) = -2.46</td>
<td>ADF(1) = -3.12 **</td>
<td>P = 0.277</td>
<td>P = 0.378</td>
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<tr>
<td>France</td>
<td>No</td>
<td>Yes</td>
<td>Yes</td>
<td>Yes</td>
<td>No (d)</td>
<td>Yes</td>
<td>Fails conditions ii) and iii)</td>
</tr>
<tr>
<td></td>
<td>ADF(1) = 0.52</td>
<td>ADF(1) = -2.73 *</td>
<td>P = 0.079</td>
<td>P = 0.824</td>
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<tr>
<td>Italy</td>
<td>No</td>
<td>Yes</td>
<td>No</td>
<td>---</td>
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<td>---</td>
<td>Fails condition i)</td>
</tr>
<tr>
<td></td>
<td>ADF(1) = -0.28</td>
<td>ADF(1) = -3.62 ***</td>
<td>P = 0.033</td>
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<tr>
<td>Spain</td>
<td>No</td>
<td>Yes</td>
<td>Yes</td>
<td>Yes</td>
<td>Yes</td>
<td>Yes</td>
<td>Fails condition ii)</td>
</tr>
<tr>
<td></td>
<td>ADF(1) = -0.53</td>
<td>ADF(1) = -2.93 **</td>
<td>P = 0.099</td>
<td>P = 0.287</td>
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</tr>
<tr>
<td>Portugal</td>
<td>No</td>
<td>No</td>
<td>---</td>
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<td>---</td>
<td>---</td>
<td>Fails condition i)</td>
</tr>
<tr>
<td></td>
<td>ADF(1) = -0.39</td>
<td>ADF(1) = -1.91</td>
<td></td>
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</tr>
</tbody>
</table>

(a) The significance level of the ADF tests is marked: * for 10%, ** for 5%, *** for 1%. In all other tests, P stands for the corresponding p-value.
(b) We have $\lambda=0$ in the model with constant term, with P=0.064
(c) We have $\lambda=\theta_1=\ldots=\theta_r=0$ in the model with constant term, with P=0.071
(d) We have $\lambda=0$ in the model with constant term, with P=0.131
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