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WP 9-01

September 2001

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***USING THE FIRST PRINCIPAL COMPONENT
AS A CORE INFLATION INDICATOR***

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(September, 2001)

ABSTRACT

This paper investigates the consequences of non-stationarity for the principal components analysis and suggests a data transformation that allows obtaining smoother series for the first principal component to be used as a core inflation indicator. The paper also introduces a theoretical model, which allows interpreting core inflation as a common stochastic trend to the year-on-year rates of change of the price indices of the basic CPI items. Finally, it is shown that the first principal component computed in real time meets the evaluation criteria introduced in Marques et al. (2000).

1. INTRODUCTION

Coimbra and Neves (1997) introduced a new a core inflation indicator based on the principal components approach. The Banco de Portugal has used such indicator, which more specifically corresponds to the first principal component, to analyse price developments, together with other core inflation measures, such as trimmed means. This new indicator, based on the principal components approach, has proved to exhibit some nice properties when evaluated against the conditions proposed in Marques *et al.* (1999, 2000).

The aim of this study is twofold. First, it investigates the consequences of non-stationarity for the computation of principal components. In fact, this technique was initially developed under the assumption of stationary variables. However, this is not the case for the large bulk of the year-on-year rate of change of prices indices pertaining to the basic items of the Consumer Price Index (CPI). Second, it tests in a more thorough way than in Marques *et al.* (1999, 2000) the first principal component against the general conditions required for a core inflation indicator. In fact, in those studies the indicator analysed was computed using all the available sample information and not, as it should, using only the information available up to and including the corresponding month. This is important because, in practice, we have to use the indicator computed in real time, and so it matters whether those conditions are still met under these circumstances.

Additionally, this study also presents a theoretical model that allows interpreting core inflation as a common stochastic trend for the year-on-year rates of change of the price indices of the basic items included in the CPI.

The first principal component computed taking into account the two above mentioned aspects, that is, both the consequences of non-stationarity and of using information available only up to and including the corresponding month, meets all the proposed conditions for a core inflation indicator. Furthermore, it is slightly less volatile than the current version of the first principal component that has been computed by the Banco de Portugal for some years now. Thus this new indicator appears to be an additional useful tool to be used in the analysis of price developments in Portugal.

This paper is organised as follows. Section 2 discusses the principal components technique and describes the main methodological changes introduced in order to account for non-stationarity. Section 3 presents and analyses a theoretical model for core inflation in the principal components framework. Section 4 analyses the properties of the indicator against the criteria introduced in Marques *et al.* (2000) and section 5 summarises the main conclusions.

2. *PRINCIPAL COMPONENTS ANALYSIS*

The principal components analysis is a statistical technique that transforms the original set of, say, N variables π_i , into a smaller set of linear combinations that account for most of the variance of the original set. For example, in our case, π_i can be thought of as the year-on-year rate of change of the i^{th} basic item included in the CPI.

It is well known that principal components analysis is not scale invariant. This is why it is customary to previously standardise the original series in order to get comparable data and then proceed with the principal components analysis on the transformed data.

Let x_{it} stand for the standardised π_i variable. By definition we have

$$x_{it} = \frac{\pi_{it} - \bar{\pi}_i}{s_i} \quad (1)$$

where $\bar{\pi}_i$ is the sample mean of π_i and s_i the corresponding standard-error. Now if X denotes the (TxN) matrix where T is the number of observations (sample period) and N is the number of standardised variables we may write

$$X = \begin{bmatrix} x_{11} & \cdots & x_{N1} \\ \vdots & \ddots & \vdots \\ x_{1T} & \cdots & x_{NT} \end{bmatrix} \quad (2)$$

As we shall see below this standardisation is generally a sensible transformation of the data, but there are other possibilities. In practice the transformation to be performed on the original data depends on the very nature of the data (statistical properties) as well as on the purposes of the analysis.

Let us assume, for the time being, that X is the matrix with the standardised variables as defined in (1). The principal components analysis aims at finding a new set of variables obtained as linear combinations of the columns of the X matrix, which are orthogonal to each other, and are such that the first accounts for the largest amount of the total variation in the data, the second for the second largest amount of the variation in the data not already accounted for by the first principal component, and so on and so forth. If we let z_{1t} denote the first of these new variables, we may write

$$z_{1t} = \beta_{11}x_{1t} + \beta_{21}x_{2t} + \dots + \beta_{N1}x_{Nt} \quad t = 1, 2, \dots, T \quad (3)$$

or in matrix form $Z_1 = X\beta_1$. The sum of squares of Z_1 is given by $Z_1'Z_1 = \beta_1'X'X\beta_1$ and the purpose of the analysis is to find out the β_1 vector that maximises $Z_1'Z_1$, subject to the restriction $\beta_1'\beta_1 = 1$, that is, to solve the problem:

$$\begin{aligned} \text{Max: } Z_1'Z_1 &= \beta_1'R\beta_1 \\ \text{s.t. } \beta_1'\beta_1 &= 1 \end{aligned} \quad (4)$$

where $R = X'X$. The condition $\beta_1'\beta_1 = 1$ is an identifying restriction that forces a finite solution for the maximum of $Z_1'Z_1$. Otherwise, just by re-scaling the β_1 vector it would be possible to arbitrarily increase the variance of the first principal component. The $R = X'X$ matrix is usually referred to as the input matrix, and if it happens that the entries in the X are the standardised variables as in (2), then R is the sampling correlation coefficients matrix for the π_{it} variables, that is,

$$R = [r_{ij}] = \sum_{t=1}^T x_{it}x_{jt} = \frac{\sum_{t=1}^T (\pi_{it} - \bar{\pi}_i)(\pi_{jt} - \bar{\pi}_j)}{s_i s_j} \quad (5)$$

Alternatively, the matrix $R = X'X$ can be written as

$$R = X'X = D_s^{-\frac{1}{2}} S D_s^{-\frac{1}{2}} \quad (6)$$

where

$$S = [s_{ij}^2] = \sum_{t=1}^T (\pi_{it} - \bar{\pi}_i)(\pi_{jt} - \bar{\pi}_j), \quad D_s = \text{diag}[s_{jj}^2] \quad \text{and} \quad D_s^{-\frac{1}{2}} = \text{diag}\left[\frac{1}{\sqrt{s_{jj}^2}}\right] = \text{diag}\left[\frac{1}{s_{jj}}\right] \quad (7)$$

Note that S is the variance-covariance matrix of the year-on-year rate of change of price indices of the basic CPI items, and D_s is the diagonal matrix of the corresponding variances (the main diagonal of S).

One can show that the solution for problem (4) is obtained by taking β_1 equal to the normalised eigenvector corresponding to the largest eigenvalue of the $R = X'X$ matrix. In other words, the optimal β_1 , say $\hat{\beta}_1$, is given by the solution for the homogeneous system

$$(R - \lambda_1 I)\hat{\beta}_1 = 0 \quad (8)$$

where λ_1 stands for the largest eigenvalue of R . Similarly, the solution for the second principal component is obtained by making β_1 equal to the normalised eigenvector corresponding to the second largest eigenvalue of $R = X'X$ and so on and so forth.¹

If we let Z_1^* denote the first principal component computed using $\hat{\beta}_1$ we have by definition:

$$Z_1^* = X\hat{\beta}_1 = (YD_s^{-\frac{1}{2}})\hat{\beta}_1 \quad (9)$$

where Y stands for the matrix of the centred year-on-year rate of change of price indices $(\pi_{it} - \bar{\pi}_i)$ of the basic CPI items. This first principal component can be obtained equivalently as $Z_1^* = Y\hat{\gamma}_1$, where $\hat{\gamma}_1$ solves the problem

$$Max. Z_1'Z_1 = \frac{\gamma_1'S\gamma_1}{\gamma_1'D_s\gamma_1} \quad (10)$$

It can be shown that $\hat{\gamma}_1$ is given by the solution to the system

$$(S - \lambda_1^* D_s)\hat{\gamma}_1 = 0 \quad (11)$$

where λ_1^* now represents the largest eigenvalue of S relatively to D_s (or equivalently the largest eigenvalue of $D_s^{-1}S$)².

¹ A proof of this result can be found, for instance, in Johnston (1984).

² See, for instance Carroll and Green (1997)

This formulation of the problem as a penalised likelihood suggests an alternative interpretation of the standardisation. Noting that $\gamma_1' S \gamma_1$ in (10) stands for the variance of $Z_1 = Y\gamma_1$, we can conclude that using R as an *input* in (4) amounts to determining the linear combination which has the largest variance relatively to the variance that would arise if the individual series were uncorrelated. But this formulation also suggests that the normalisation implied by the use of R as *input* might not be the only one, and that the choice can be adapted to the problem at hand. For instance, sometimes the X matrix is defined with entries $x_{it} = (\pi_{it} - \bar{\pi}_i)$, i.e. with variables subtracted from their means. In this case, the input matrix $R = X'X$ is the variance-covariance matrix of the original data. The use of the variance-covariance matrix as the input matrix could be acceptable if the original variables do exhibit variances that do not differ much among them. Otherwise the first principal component tends to be dominated by the variables with the largest variances. As the variance is scale dependent the solution to such a case is exactly to use standardised variables. See, for instance, Dillon and Goldstein (1984).

The principal components analysis was first developed under the assumption of stationary variables. In case of stationarity standardisation has an immediate statistical interpretation. However, in the Portuguese case, it is possible to show that the year-on-year rates of price changes of most basic CPI items behave as non-stationary variables. Particularly, for most of these series the null of a unit root is not rejected. In such a case, two different questions arise quite naturally. On the one hand the issue of whether the principal components analysis still applies for variables integrated of order one and, on the other, whether the classical standardisation is still to be used given the purpose of building a trend inflation indicator.

The answer to the first question is yes. The principal components analysis is still applicable with non-stationary variables. The so-called principal components estimator with non-stationary variables was first utilised by Stock and Watson (1988). Recently, Harris (1997) showed that this estimator could be used to estimate cointegrating vectors. In this context, the estimator for β_1 that allows minimising the variance of z_{1t} and so obtaining the cointegrating vector that stationarises z_{1t} in (3) is given by the eigenvector corresponding to the smallest eigenvalue of $X'X$. Harris (1997) demonstrated that the estimator for β_1 is super-consistent be it an estimator of a cointegrating vector or an estimator of a principal component.³ However, it is in general asymptotically inefficient, which lead the author to develop a modified principal

³ See also Hall *et al.* (1999). The authors also discuss the case in which the matrix X includes not only $I(1)$ variables but also $I(0)$ variables, showing that, even then, the principal components estimator is super-consistent.

components estimator that is asymptotically efficient for a wide range of data generating processes. In our case, the large number of variables (basic CPI items) would make the implementation of the correction suggested in Harris (1997) very demanding, and the comparatively small number of observations would reduce, perhaps substantially, the potential efficiency gain. For this reason, the correction was not introduced in the estimator for this study.

Let us now answer the second question. The estimated coefficients of the β_1 vector in (3) can be seen as representing the contribution (weight) of each basic item for the definition of the first principal component. Since we aim at maximising the variance of z_{1t} in (3) the corresponding estimator will attach a larger weight to the components with a larger variance. The common standardisation, which is obtained by subtracting the mean and dividing by the standard error, is adequate when the original variables are stationary. However, when variables are integrated of order one the sampling variance is the larger the larger the change in the average level of the variable during the sample period. Thus, the series exhibiting strong increasing or decreasing trends in the sample will appear as very volatile no matter how smooth they are. In other words, in the case of integrated variables the empirical variance is not a good measure of volatility.

If the purpose is to obtain a core inflation indicator then we should care about the degree of smoothness of the first principal component and thus to look for linear combinations of the year-on-year variation rates of the basic CPI items with a large signal (variance) and not too much volatility. Define the smoothness of an integrated variable as the variance of the first differences, and let V be the variance-covariance matrix of the year-on-year variation rates of the basic CPI items (the columns in Y). Then, this purpose can be formalised, in an analogous way to (10), as

$$Max. Z_1' Z_1 = \frac{\gamma_1' S \gamma_1}{\gamma_1' D_v \gamma_1} \quad (12)$$

where D_v is the diagonal matrix containing the main diagonal of V . The solution of (12) satisfies

$$(S - \lambda_1^* D_v) \hat{\gamma}_1 = 0 \quad (13)$$

where λ_1^* is now the largest eigenvalue of S relative to D_v , and $\hat{\gamma}_1$ is the eigenvector associated to λ_1^* . Again, the first principal component will be given by $Z_1^* = Y \hat{\gamma}_1$.

The conditions in (13) can also be written as⁴

$$(D_v^{-\frac{1}{2}} S D_v^{-\frac{1}{2}} - \lambda_1^* I) D_v^{\frac{1}{2}} \hat{\gamma}_1 = 0 \quad (14)$$

Thus, it is clear that Z_1^* could also be obtained by solving

$$\begin{aligned} \text{Max. } Z_1' Z_1 &= \beta_1' D_v^{-\frac{1}{2}} S D_v^{-\frac{1}{2}} \beta_1 \\ \text{s.t. } \beta_1' \beta_1 &= 1 \end{aligned} \quad (15)$$

and applying the resulting estimated parameter vector $\hat{\beta}_1^*$ to the matrix of the centred year-on-year rate of change of the basic CPI items (Y), previously standardised by the standard deviation of the differences, $Z_1^* = (Y D_v^{-\frac{1}{2}}) \hat{\beta}_1^*$. In other words, this amounts to applying the principal components analysis method, taking in (1)

$$x_{it} = \frac{\pi_{it} - \bar{\pi}_i}{\sigma_{\Delta i}} \quad (16)$$

where π_{it} denotes the year-on-year rate of change of the i^{th} basic CPI item, $\bar{\pi}_i$ the corresponding sample mean e $\sigma_{\Delta i}$ the standard error of $\Delta \pi_{it}$.

At last, it is also important to address two additional questions that have consequences on the way the indicator is computed, i.e. the need to be computable in real time and to be re-scaled.

It is usually required that a core inflation indicator should be computable in real time.⁵ The way to solve this problem is to build a series of first estimates of z_{1t} . In other words the indicator based on the principal components analysis was constructed by picking up, for each period t , the figure for the principal component we obtain from (3) by including in the X matrix only the observations available up to period t . Of course, this process can only be used after allowing for a long enough period used to compute the first estimate. In our case, given that the

⁴ Pre-multiplying (13) by $D_v^{-1/2}$ we get $(D_v^{-\frac{1}{2}} S - \lambda_1^* D_v^{\frac{1}{2}}) \hat{\gamma}_1 = 0$, but

$$(D_v^{-\frac{1}{2}} S - \lambda_1^* D_v^{\frac{1}{2}}) \hat{\gamma}_1 = (D_v^{-\frac{1}{2}} S D_v^{-\frac{1}{2}} D_v^{\frac{1}{2}} - \lambda_1^* D_v^{\frac{1}{2}}) \hat{\gamma}_1 = (D_v^{-\frac{1}{2}} S D_v^{-\frac{1}{2}} - \lambda_1^* I) D_v^{\frac{1}{2}} \hat{\gamma}_1$$

⁵ See, for instance, Marques *et al.* (2000)

sample is very short we decided, for the purpose of analysing the properties of the corresponding indicator, in the terms of section 4, to retain the initial figures even though, in rigor, they are not first estimates. This way, for the period 1993/7 –1997/12 the indicator is made up of estimates obtained using the data up to 1997/12 and after that it is in fact made up of first estimates computed as explained above. One must notice that this new indicator allows a more rigorous analysis of the first principal component indicator than the one evaluated in Marques *et al.* (1999, 2000).

Let us now address the re-scaling issue. The average level of the principal component in (3), being obtained after “standardising” the original data, is not comparable to the inflation average level during the sample period. To be used as a core inflation indicator it has to be re-scaled so that the two series may exhibit the same average level. Even though there are several alternative procedures the easiest one to implement is to run a regression equation between the inflation rate and the first principal component and to define the re-scaled indicator as the one corresponding to the fitted values of the regression.⁶ In our case in order to get an estimator computable in real time, we have decided to estimate successive regressions each time including an additional observation.

The analysis of this indicator made up of first estimates, which we shall denote as PC1 is carried out in section 4. For comparability reasons an indicator, also computed in real time after 1998/1, was constructed, in which the conventional standardisation was performed.⁷ This indicator shall be denoted below as PC2.

3. A THEORETICAL MODEL FOR THE TREND OF INFLATION

In this section we show as the principal components analysis may be used to derive a consistent estimate for the trend of inflation. Let us assume that the price change of the i^{th} CPI item can be decomposed as the sum of two distinct components. The first that we shall call the permanent component whose time profile is basically determined by the trend of inflation and the second usually referred to as the temporary component, which basically is the result of the idiosyncratic shocks, specific to the market of the i^{th} good. In generic form we write

$$\pi_{it} = a_i + b_i \pi_t^* + \varepsilon_{it}; \quad i = 1, \dots, N; \quad t = 1, \dots, T; \quad (17)$$

⁶ This was the methodology used, for instance, in Coimbra and Neves (1997).

⁷ That is, using the standard error of π_{it} and not of $\Delta\pi_{it}$.

where π_{it} , once again, stands for year-on-year price change of the i^{th} item, π_t^* for the trend of inflation and ε_{it} for the temporary component.

Assuming that the π_{it} variables are integrated of order one, it follows that π_t^* is also integrated of order one. In turn, each ε_{it} is, by construction, a zero-mean stationary variable. Thus, equation (7) posits a cointegrating relationship between the change of prices of the i^{th} item and the trend of inflation.⁸ We assume at this disaggregation level that there are some CPI items whose price changes, even though determined in the long run by the trend of inflation, do not necessarily exhibit a parallel evolution vis-à-vis the trend of inflation (so that we can have both $a_i \neq 0$ and $b_i \neq 1$).

One should notice that the general formulation suggested in (7) where we may have $a_i \neq 0$ and $b_i \neq 1$ is not incompatible with the usual hypothesis made in the literature, at the aggregate level, which decomposes the economy-wide inflation rate as the sum of the trend of inflation and a transitory component

$$\pi_t = \pi_t^* + u_t . \quad (18)$$

To see that let us start by noticing that the inflation rate measured by the year-on-year CPI rate of change may be written as $\pi_t = \sum_{i=1}^N w_{it} \pi_{it}$, with $w_{it} = \alpha_i \frac{P_{i,t-12}}{P_{t-12}}$, where α_i represents the (fixed) weight of the i^{th} item in the CPI, P_{it} the corresponding price index and P_t the CPI itself. Notice also that we have $\sum_{i=1}^N w_{it} = 1$, even though the w_{it} are time varying.

If you multiply the N equations (7) by the w_{it} weights we get

$$\sum_{i=1}^N w_{it} \pi_{it} = \sum_{i=1}^N w_{it} a_i + \sum_{i=1}^N w_{it} b_i \cdot \pi_t^* + \sum_{i=1}^N w_{it} \varepsilon_{it} \quad (19)$$

that is

$$\pi_t = \phi_{0t} + \phi_{1t} \pi_t^* + v_t \quad (20)$$

⁸ Notice however that the method is also applicable even if some π_{it} are stationary, i.e. if some b_i are zero [see Hall *et al.* (1999)].

Now if we have in (10)

$$\begin{cases} E[\phi_{0t}] = E\left[\sum_{i=1}^N w_{it} a_i\right] = 0 \\ E[\phi_{1t}] = E\left[\sum_{i=1}^N w_{it} b_i\right] = 1 \end{cases} \quad (21)$$

the relation suggested in (8) will be satisfied.⁹

We will now show how the principal components method can be used to consistently estimate π_t^* in the context of model (17). This means that the first principal component can be thought of as a common stochastic trend to the year-on-year rate of change of each basic CPI item. Let

$$x_{it} = \frac{\pi_{it} - \bar{\pi}_i}{\gamma_i} \text{ and } f_t = \pi_t^* - \bar{\pi}^* \quad (22)$$

where $\bar{\pi}_i$ and $\bar{\pi}^*$ are the sample means of π_{it} and π_t^* , respectively, while γ_i and N are non-zero parameters. Noting that $\bar{\pi}_i = a_i + b_i \bar{\pi}^* + \bar{\varepsilon}_i$, (17) can be rewritten as

$$x_{it} = \beta_i f_t + \mu_{it} \quad i = 1, \dots, N; t = 1, \dots, T; \quad (23)$$

where $\beta_i = b_i / \gamma_i$ and $\mu_{it} = \frac{\varepsilon_{it} - \bar{\varepsilon}_i}{\gamma_i}$ are stationary with zero mean. In this equation, f_t is usually called the common (stochastic) trend since, in the long run, this is the variable that determines the behaviour of the x_{it} .

The principal components analysis allows us to obtain a super-consistent estimate for f_t . As we saw in the previous section, the principal components are computed from the cross-products matrix,

$$R = X'X \quad (24)$$

⁹ Note that for (21) to hold when the w_{it} follow I(1) processes, it must be true that $a = (a_1, \dots, a_N)$ e $b = (b_1, \dots, b_N)$ are cointegrating vectors for the w_{it} variables.

and consist on N linear combinations of x_{it} , weighted by the eigenvectors of $X'X$. The N equations in (23) can be written as a system:

$$\left\{ \begin{array}{l} x_{1t} = \beta_1 f_t + \mu_{1t} \\ \vdots \\ x_{it} = \beta_i f_t + \mu_{it} \\ \vdots \\ x_{N-1,t} = \beta_{N-1} f_t + \mu_{N-1,t} \\ x_{Nt} = \beta_N f_t + \mu_{Nt} \end{array} \right\} \Rightarrow \left\{ \begin{array}{l} x_{1t} = \frac{\beta_1}{\beta_N} x_{Nt} + \left(\mu_{1t} - \frac{\beta_1}{\beta_N} \mu_{Nt} \right) \\ \vdots \\ x_{it} = \frac{\beta_i}{\beta_N} x_{Nt} + \left(\mu_{it} - \frac{\beta_i}{\beta_N} \mu_{Nt} \right) \\ \vdots \\ x_{N-1,t} = \frac{\beta_{N-1}}{\beta_N} x_{Nt} + \left(\mu_{N-1,t} - \frac{\beta_{N-1}}{\beta_N} \mu_{Nt} \right) \end{array} \right. \quad (25)$$

that is, there are $N-1$ stationary and independent linear combinations of x_{it} , which can take the form

$$v_{it} = x_{it} - \frac{\beta_i}{\beta_N} x_{Nt}; \quad i = 1, \dots, N-1 \quad (26)$$

In this case, the principal components associated with the $N-1$ smaller eigenvalues are stationary, and the corresponding eigenvectors estimate a base for the space of cointegrating vectors. From the equations in (26), one can immediately check that one possible base for the space of cointegrating vectors is given by

$$B = \begin{bmatrix} 1 & 0 & \dots & 0 \\ 0 & 1 & \dots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & \dots & 1 \\ -\frac{\beta_1}{\beta_N} & -\frac{\beta_2}{\beta_N} & \dots & -\frac{\beta_{N-1}}{\beta_N} \end{bmatrix} \quad (27)$$

On the other hand, the principal component associated to the largest eigenvalue is $I(1)$, since it is the linear combination of x_{it} with maximum variance. Harris (1997) has shown that this principal component is a super-consistent estimator of f_t (up to a scale factor). Since the eigenvector in question is orthogonal to the space defined by B , it must be of the form $k\alpha$, with $k \neq 0$ and

$$\alpha = (\beta_1, \beta_2, \dots, \beta_N) = \left(\frac{b_1}{\gamma_1}, \frac{b_2}{\gamma_2}, \dots, \frac{b_N}{\gamma_N} \right) \quad (28)$$

and so

$$f_t = k \sum_{i=1}^N \beta_i x_{it} = k \sum_{i=1}^N \frac{b_i}{\gamma_i} x_{it} \quad (29)$$

From this estimate for f_t , one can obtain an estimate for π_t^* through a linear transformation, by applying the least squares method to

$$\pi_t = \delta_0 + \delta_1 f_t + \eta_t \quad (30)$$

This way, one guarantees that π_t^* has the same mean level as inflation, during the sample period.

One can note that instead of (17), the literature on core inflation indicators usually considers the relation

$$\pi_{it} = \pi_t^* + \varepsilon_{it} \quad (31)$$

i.e., the restrictions $a_i = 0$ and $b_i = 1$, are imposed *a priori*, for all i . Notice that (31) postulates that in each market, the inflation rate is essentially equal to the sum of core inflation (or the mean inflation rate of the economy) and a specific component which accommodates relative prices changes.

If we start from equation (31), then equation (23) holds with $\beta_i = 1/\gamma_i$, reducing the eigenvector in (28) to

$$\alpha = \left(\frac{1}{\gamma_1}, \frac{1}{\gamma_2}, \dots, \frac{1}{\gamma_N} \right) \quad (32)$$

Considering that these weights are associated to the x_{it} variables defined in (22), the weights associated with the original π_{it} are proportional to

$$\varphi = \left(\frac{1}{\gamma_1^2}, \frac{1}{\gamma_2^2}, \dots, \frac{1}{\gamma_N^2} \right) \quad (33)$$

Thus, the weights only depend on the measure used to normalise the individual items, and the principal components analysis is unnecessary (since there is nothing to be estimated). Note that,

taking γ_i in (22) as the standard deviation of $(\pi_{it} - \pi_t)$, the resulting “first principal component” is identical to the so-called “neo-Edgeworthian” index, where the weights are proportional to the inverses of the variances of $(\pi_{it} - \pi_t)$.¹⁰

4. ANALYSING THE PROPERTIES OF THE INDICATOR

In this section the properties of the two indicators PC1 e PC2 described in section 2 are evaluated. The evaluation of the trend inflation indicators follows the criteria proposed in Marques *et al.* (1999, 2000). Remember that these criteria are the following:

- i) the difference between observed inflation and the trend indicator must be a zero- mean stationary variable;
- ii) the trend indicator must behave as an attractor for the rate of inflation, in the sense that it provides a leading indicator of inflation;
- iii) the observed inflation should not be an attractor for the trend inflation indicator.

To test these conditions we may proceed in different ways. The verification of condition i) may be carried out by testing for cointegration in the regression equation $\pi_t = \alpha + \beta\pi_t^* + u_t$, with $\beta = 1$ and $\alpha = 0$, where π_t stands for the year-on-year inflation rate and π_t^* for the trend inflation indicator. In turn, this test can be implemented in two steps. First run the unit root test on the series $d_t = (\pi_t - \pi_t^*)$ with a view to show that d_t is a stationary variable. Second, test the null hypothesis $\alpha = 0$, given that d_t is stationary.

To test the second and third conditions we need to specify dynamic models for both π_t and π_t^* . For the technical details the reader is referred to Marques *et al.* (2000).

Both the PC1 and PC2 indicators meet the three suggested conditions. We note that, by construction, we should expect both indicators to be unbiased estimators, that is, to meet the second part of condition i).

Figure 1 shows that both indicators behave very much like what we would expect from a core inflation indicator. Namely, CP1 and CP2 are smoother than inflation, and tend to be higher than inflation when this is low and to be below inflation when this is particularly high. Furthermore, under these circumstances, we see that it is the inflation that converges to the indicator and not the other way around. Figure 1 also shows that CP2 is slightly more volatile

¹⁰ For a detailed description of the “neo-Edgeworthian” index see, e.g., Marques *et al.* (2000), pp. 9-10.

than CP1¹¹, so that the theoretical advantages put forward in the previous section, become now apparent.

Let us now compare the weights in the CPI with the corresponding weights in the first principal component, for the different items.

Figure 2 depicts the relation between the weights of each CPI item in the CP1 indicator and the corresponding volatility (evaluated by $\sigma_{\Delta i}$, the standard error of the first differences), both computed with the data available for the whole sample period. It turns out that all the items with a significant weight exhibit a relatively low volatility and that the items with larger volatility have weights close to zero. It thus exists a negative relationship between the weights and the volatility for each item. On the contrary, as we can see in Figure 3, there is no significant relationship between the weight of each item in the CPI index and the corresponding weight in the first principal component.

Figure 4 depicts the CPI weights of 9 CPI aggregates and the corresponding weights in the first principal component.¹² The first two aggregates are basically composed of the items excluded from the traditional “excluding food and energy” indicator. The remaining aggregates are the same as in the CPI. It turns out that the weights of the aggregates “unprocessed food” and “energy” in the first principal component are smaller than their weights in the CPI. This is also true, even though to a lesser extent, for “processed food” and “Transportation and Communications (excluding energy)”. All the remaining aggregates exhibit a larger weight in the first principal component than in the CPI.

Summing up we may conclude that the most volatile series reduce their weights in the first principal component vis-à-vis the CPI, and vice-versa for the smoothest series. This fact explains why the CP1 indicator is smoother than the observed rate of inflation.

Finally it is important to note that the CP1 indicator, even though it seems to behave rather satisfactorily under normal circumstances, it may nevertheless exhibit stability problems under special circumstances, namely if a change in the number, in the definition or in data collecting process of the basic CPI items occurs. In this case the use of the first principal

¹¹ The standard error of the PC1 first differences is 0.093 p.p. and the one of PC2 é 0.120 p.p.. Both these standard errors are significantly lower than the one of the first differences of observed rate of inflation, which is 0.297 p.p. .

¹² The estimated weights of some basic items in the first principal component appear with a negative sign. However, most of them appear not to be significantly different from zero and their accumulated weight is rather small (about -1.86%). For this reason we choose to keep them in the figure. We note that the weight of the aggregate “unprocessed food”, the most affected by this problem, will be 5.32% instead of 3.78% if those negative weight have been removed.

component should be complemented with more robust indicators such as some limited influence estimators currently used by the Banco de Portugal.

5. CONCLUDING REMARKS

In this paper we re-estimate and re-evaluate the first principal component as trend inflation indicator. The re-estimation is done so that the indicator is computed in real time and re-evaluation is carried out after allowing for the presence of a unit root in the generation processes of the price changes series.

The new indicator meets all the properties required for a core inflation indicator. On the one hand it turns out that only the relatively smooth series exhibit significant weights in first principal component, the weights of the volatile series being almost null. In particular the weight of the volatile aggregates “unprocessed food” and “energy” is much smaller in the first principal component than in the consumer price index. This is why the core inflation indicator is much less volatile than recorded inflation. On the other hand recorded inflation tends to converge for the first principal component whenever there is a significant difference between them.

We thus think that this new core inflation indicator may play a useful role in the analysis of price developments in the Portuguese Economy.

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Figure 1

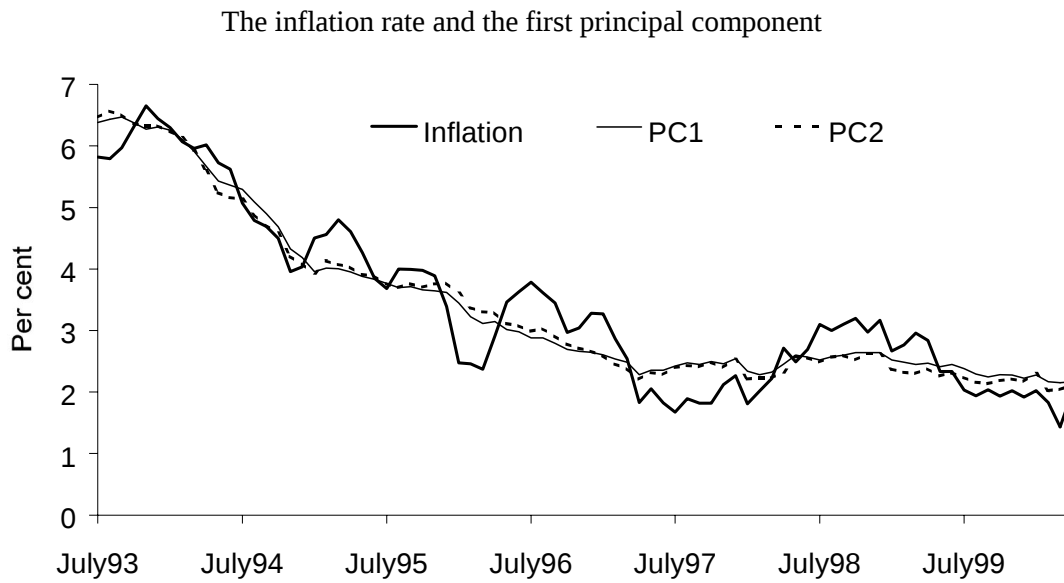


Figure 2

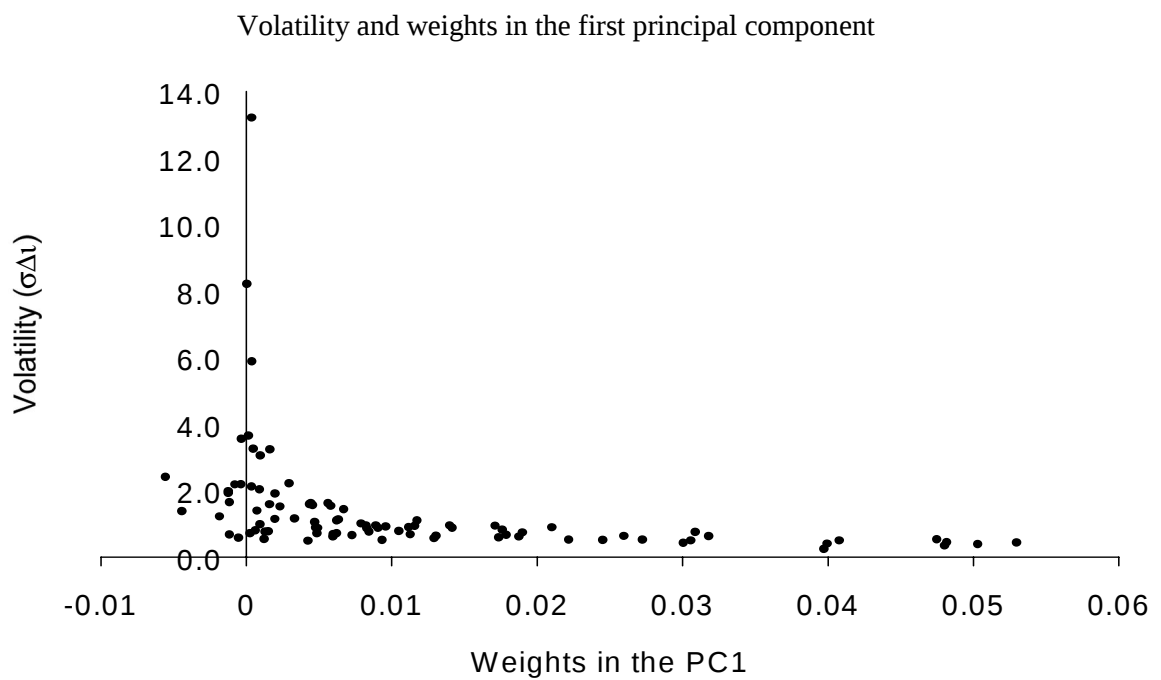


Figure 3
Weights in the CPI and in the first principal component

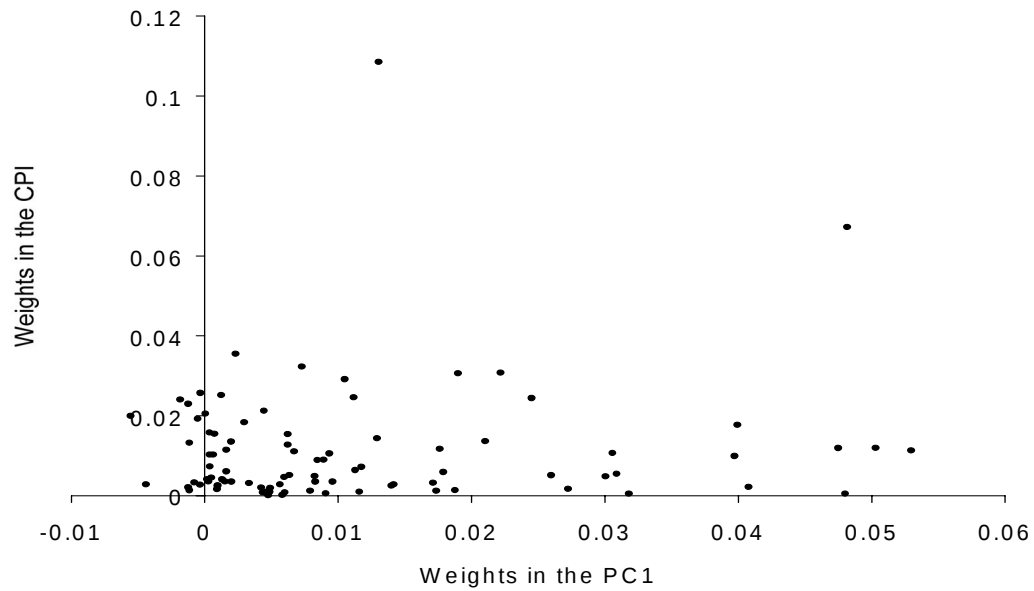
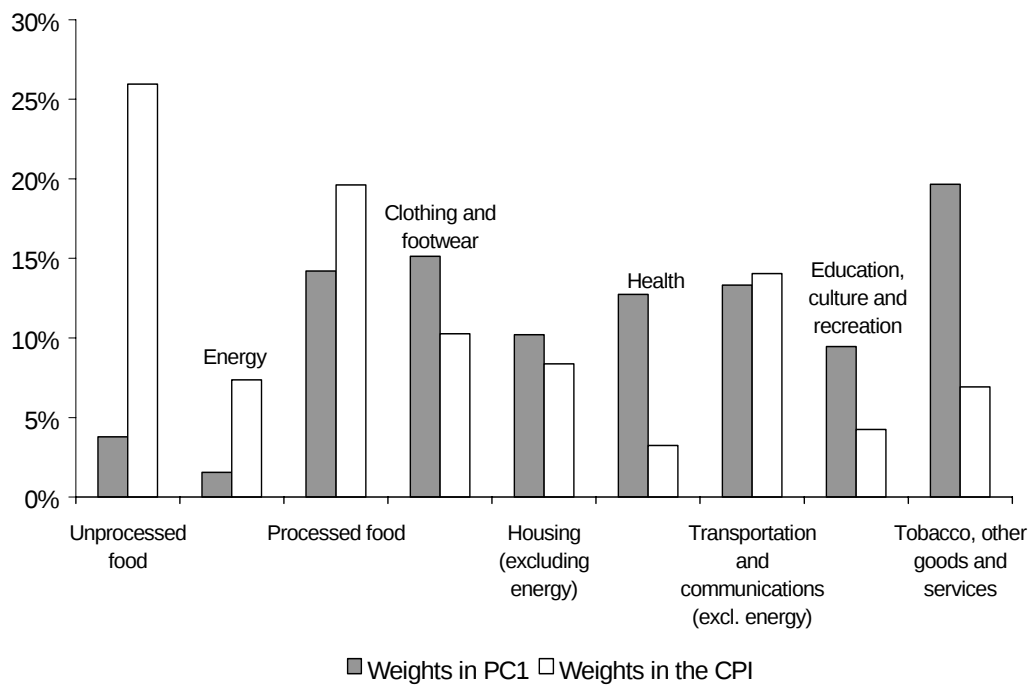


Figure 4
Weights of some aggregates in the CPI and in the first principal component



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