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Av. Universitaria 1801, Lima 32 – Perú.
Teléfono: (51-1) 626-2000 anexos 4950 - 4951
Fax: (51-1) 626-2874
econo@pucp.edu.pe
www.pucp.edu.pe/departamento/economia/

Encargado de la Serie: Luis García Núñez
Departamento de Economía – Pontificia Universidad Católica del Perú,
lgarcia@pucp.edu.pe

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A Factorial Decomposition of Inflation in Peru. An Alternative Measure of Core Inflation

Alberto Humala
Central Bank of Peru

Gabriel Rodríguez
Pontificia Universidad Católica del Perú

Abstract

A dynamic factorial decomposition model of inflation is estimated using Peruvian monthly data for 1995:01-2008:07. This model allows identification of changes in three relevant inflation components: idiosyncratic relative prices, aggregate relative prices, and absolute prices. Furthermore, following Reis and Watson (2007), the model allows measuring pure inflation as the common factor in the inflation rate that has a proportionate effect to all prices and that is not correlated with relative price changes at any period of time. This pure inflation estimate relates closely to standard measures of core inflation. Results are robust to different lag structures and various stochastic assumptions on the estimated factors.

Keywords: Factoral Decomposition, Pure Inflation, Core Inflation, Price Changes.

JEL: C32, C43, E31.

Resumen

Un modelo de descomposición factorial dinámico es estimado usando datos mensuales para Perú para el periodo 1995:01-2008:07. Este modelo permite la identificación de cambios en tres importantes componentes de la inflación: precios relativos idiosincráticos, precios relativos agregados y precios absolutos. Asimismo, siguiendo la metodología de Reis y Watson (2007), el modelo permite encontrar una medida de inflación pura como el factor común en la tasa de inflación que tiene un efecto proporcional a todos los precios y que no está correlacionado con cambios en precios relativos en cualquier periodo del tiempo. Este estimado de inflación pura está vinculado de manera cercana con otras medidas frecuentemente utilizadas de inflación subyacente. Los resultados son robustos a diferentes estructuras de rezagos y a varios supuestos sobre la (no) estacionariedad de los factores estimados.

Palabras Claves: Descomposición Factorial, Inflación Pura, Inflación Subyacente, Cambio en Precios

Classificación JEL: C32, C43, E31.

A Factorial Decomposition of Inflation in Peru. An Alternative Measure of Core Inflation[1]

Alberto Humala[2]
Central Bank of Peru

Gabriel Rodríguez[3]
Pontificia Universidad Católica del Perú

1 Introduction

The primary goal of factor analysis is to account for the presence of correlation among a group of variables. Thus, for instance, cross-sectional correlation of asset returns may be explained by a single factor, according to capital asset pricing theory. Alternatively, correlation among many macroeconomic variables could be due to some common shocks. In these examples, a common feature is that a large number of variables are linked with a small number of unobserved components that give rise to cross correlations.

In respect to this, Bai (2004) offers an excellent survey and Reis and Watson (2007) use a dynamic factor model for quarterly changes in consumption good prices to decompose them into three price variations: idiosyncratic relative prices, aggregate relative prices, and absolute prices. Furthermore, Reis and Watson (2007) identify a measure of "pure" inflation as the common component in the price variation rates that has equally proportionate effects on all prices and that is uncorrelated with relative price changes at all dates. In addition, they use this measure of "pure" inflation to re-examine some classic macroeconomic relationships such as the money to inflation link.

Previous to Reis and Watson (2007), there have been few attempts at separating absolute-price from relative-price variations. An exception is Bryan and Cechetti (1993), who use a dynamic factor model in a panel of 36 price series to measure the absolute-price change component. However, in order to identify and estimate the model, they impose the restriction

[1] We thank useful comments from participants to the XXVI Meeting of Economists of the Central Reserve Bank of Peru in November 2008. We are grateful to Reis and Watson for useful E-Mail communications and for providing us with their estimation codes. The views expressed herein are those of the authors and do not reflect necessarily those of their institutions.

[2] Address for Correspondence: Alberto Humala, Senior Researcher, Central Reserve Bank of Peru, Miró Quesada 441-445, Lima 1, Lima, Perú. +511-613-2785, Fax: +511-613-2516. E-Mail address: alberto.humala@bcrp.gob.pe.

[3] Professor, Department of Economics, Pontificia Universidad Catolica of Peru, Av. Universitaria 1801, Lima 32, Lima, Perú, Telephone: +511-626-2000 (4998), Fax: +511-626-2874. E-Mail Address: gabriel.rodriguez@pucp.edu.pe.

that relative prices are independent across goods.

In methodological terms, the research of Reis and Watson (2007) is related to large-scale dynamic factor models estimated by maximum likelihood. See, for instance, Quah and Sargent (1993); and Doz, Giannone, and Riechlin (2006). Another related approach is the principal components used by Bai and Ng (2002); Forni, Hallin, Lippi, and Riechlin (2000); and Stock and Watson (2002).

Using the above mentioned methods, Cristadoro, Forni, Riechlin, and Veronesi (2005) estimate a common factor model on a panel of quantity and price series. However, they are more interested in the forecasting properties of the approach. Alternatively, Amstad and Potter (2007) use another dynamic factor model to update daily the measure of the common component in price changes. Del Negro (2006) estimates a factor model using sector inflation rates and allows for a common component and for relative-price factors associated with durable, non-durable, and service good sectors. Furthermore, Altissimo, Mojon, and Zaffaroni (2006) estimate a common factor model using disaggregated Euro-area consumer price indices and use the model to measure the persistence in aggregate Euro-area inflation. It is worth to mention that the common factor in all these previous models is not a measure of pure inflation since it affects prices differently. A more related research to Reis and Watson (2007) is Boivin, Giannoni, and Mihov (2007) who extract macroeconomic shocks using many series that include prices but also real quantities.

Some of the above mentioned papers use factor models to replace large datasets (which often contain a few hundred series) with a small number of common factors that could be used then for price forecasting. In these cases the procedure allows using a large information set without the need to limit the number of degrees of freedom when constructing a model. These models usually do not impose restrictions on factor or loadings. However, other approaches (as Reis and Watson, 2007) use factor models to identify components with plausible economic interpretation. In such cases, restrictions to identify unobserved variables are imposed on the matrix of loadings.

In this paper, we follow Reis and Watson (2007) to estimate a dynamic factorial decomposition model of inflation using Peruvian monthly data for the period 1995:1-2008:07. This model allows identification of two relevant inflation components: changes in aggregate relative prices, and changes in absolute prices. The model also allows measuring pure inflation as the common factor in the inflation rate that has a proportionate effect on all prices and that is not correlated with relative price changes at any period of time. This pure inflation estimate relates closely to standard measures of core inflation. Results are robust to different lag structures and various stochastic assumptions on the estimated factors. Therefore, this paper aims at decomposing absolute from relative price changes in the CPI basket. As a by-product, the paper estimates an alternative measure of core inflation.[4]

Peru suffered from a hyperinflation process by the end of the 1980s, but it successfully stabilized its economy by mid-1990s.[5] A number of structural economic reforms were introduced during the first part of the 1990's, namely financial system liberalization (including a pension fund reform), trade openness, reinsertion in the international financial system, tax-system reform, sound and prudent monetary and fiscal policies, investments promotion and, in general, more market-oriented policies throughout the economy. Building upon new trends in macroeconomic variables by the late 1990's, Peru started to use money-aggregates targeting with explicit (but not yet binding) preferred inflation rates in 1994. By 2002 Peru formally adopted a fully-fledged inflation-targeting regime. The sample in this study starts from 1995 onwards. Thus, it includes a decreasing inflation scenario (towards one-digit levels) during the second part of the 1990s and a stable and low inflation period after the adoption of inflation targeting. The target is set on Consumer Price Index (CPI) inflation, but macroeconomic analysis by monetary authorities considers also a set of core inflation measures.

Next section briefly describes some aspects of the dynamic factor model. The third section describes the general model to estimate pure inflation. Section four discusses model estimation and presents some robustness analysis. Section five presents the empirical evidence and assesses the main estimating results. Lastly, Section six concludes.

2 The Dynamic Factor Model

In classical factor analysis the statistical theory is developed under a fixed number of variables (N). However, dynamic factor model analysis allows N and the number of observations (T) to be large. In particular, N can be much larger than T.

A factor model takes the following form:

$$Y_{it} = \lambda_i' F_t + u_{it}, \tag{1}$$

for $i = 1, 2, ..., N$; $t = 1, 2, ..., T$ and where Y_{it} is the observation on the

[4] Another application of the approach of Reis and Watson (2007) is Brzoza-Brzezina and Kotlowski (2009) where Polish data are used.

[5] For an account of inflation dynamics in Peru see Castillo, Humala, and Tuesta (2006).

variable i at period t, λ_i $(r \times 1)$ is a vector of factor loadings, F_t $(r \times 1)$ is a vector of factor processes, and u_{it} is the idiosyncratic error term. For example Y_{it} may represent the output growth rate for country i in quarter t; F_t is a vector of common shocks (technology shocks, financial crises, oil prices, and so on) that influence output; λ_i represents the impact of shocks on country i; and u_{it} is the country-specific shock. Yet, in another example Y_{it} may be the return of asset i in period t; F_t is a vector of factor returns with zero mean (risk premium adjusted); λ_i is a vector of factor loadings; and u_{it} is the idiosyncratic return. In this paper Y_{it} is the inflation rate for good i in period t; F_t is a vector of factor inflation rates; λ_i is a vector of factor loadings; and u_{it} is the idiosyncratic shock on good i.

Introducing additional notation, the factor model may be written as

$$Y_t = \Lambda F_t + u_t, \tag{2}$$

for $t = 1, 2, ..., T$.

Unlike classical factor analysis, dynamic factor model analysis proceeds under different assumptions. First, the number of variables N is assumed to be large and the limit theory is developed assuming that N and T go to infinity. In particular, N can be much larger than T. Second, both F_t and u_t may be serially correlated. Third, the covariance matrix of u_t does not need to be a diagonal matrix and, in fact, none of the off-diagonal elements needs to be zero. Thus, the number of parameters in the covariance matrix may be as large as the number of equations.

A key issue in factor model analysis is the estimation of the number of factors. The number of factors r may be consistently estimated using the information criterion approach proposed by Bai and Ng (2002). Let $\widehat{\sigma}^2(k)$ denotes the sum of squared residuals (divided by NT) when k factors are allowed, and let $S(\Lambda, F) = tr[(Y - \Lambda F')(Y - \Lambda F')]$ denotes the least squares objective function. Then, we have $\widehat{\sigma}^2(k) = S[\widehat{\Lambda}(k), \widehat{F}(k)]/NT$. Consider the following criterion:

$$IC(k) = \log \widehat{\sigma}^2(k) + kg(N, T). \tag{3}$$

If $g(N, T) \to 0$ and $\min[N, T]g(N, T) \to \infty$, then $\Pr(\widehat{k} = r) \to 1$, where k minimizes the information criterion. For example, $g(N, T) = (N + T)\log(NT)/NT$ satisfies the above condition.

There are some cases where the factor process F_t is a I(1) or integrated (vector) process such that $F_t = F_{t-1} + \eta_t$ and, thus, Y_t is not stationary. As shown by Bai (2004), if the idiosyncratic process u_{it} is I(0), then all Λ, F, and r can be consistently estimated. In this case, since Y_{it} (for all i) share the same common stochastic trends, F_t, Y_{it} are cointegrated to each other.

If the idiosyncratic process u_{it} is I(1) for all i and $u_{it} = u_{i,t-1} + \eta_{it}$, then there is no cointegration among the observable Y_{it}. Yet, the common stochastic trends are well defined and can be estimated consistently up to a rotation; a striking contrast with a fixed N spurious system.

3 The Model for Inflation

This section uses the previous dynamic factor model to decompose individual inflation rates into a small set of components. Notation is similar as before except that π_{it} (inflation rates of good i in period t) replaces Y_{it}.

The key feature of this approach is that price movements could be decomposed into absolute and relative price changes, so that a pure inflation measure (based on absolute variations) can be estimated. We follow closely the model by Reis and Watson (2007) to obtain this measure for Peru.

Consider π_{it} the change rate for an item i between time $t-1$ and t, so that π_t is a $N \times 1$ vector containing price changes for the N items in the CPI. Price co-movement is represented by the following lineal factorial model:

$$\pi_t = \Lambda F_t + u_t, \tag{4}$$

where F_t is a $(k \times 1)$ vector of the k factors that explain common sources for price changes; Λ is a $(N \times k)$ matrix of coefficients that measure the i^{th} item's price responds to shocks; and u_t is a $(N \times 1)$ vector that captures relative price variability associated to idiosyncratic sectorial events or measurement errors.[6]

Factors in F_t account for an important fraction of price changes and those changes provide information on aggregate shocks. Besides, this aggregate component could be split up into absolute price changes (an scalar a_t) and, potentially, several relative price components, a $(k-1)$ vector denoted by R_t such as that:

$$\Lambda F_t = \beta a_t + \Gamma R_t, \tag{5}$$

where β is a $(N \times 1)$ vector of 1's, since absolute prices changes affect all prices equiproportionally; Γ is a $N \times (k-1)$ matrix, indicating different

[6] This research, in line with Reis as Watson (2007), uses intensively data on prices to calculate alternative measures of inflation. Boivin, Giannoni, and Mihov (2007), for instance, use a similar model but estimate a common macroeconomic shock to several series of prices and real quantities.

effects from relative price changes. The question that arises immediately on ΛF_t is whether or not the common sources of variation could actually be decomposed.

Reis and Watson (2007) point out that there are indeed two inconveniences in this general factorial decomposition model. First, β could be absent from the space for Λ, meaning that there are no changes in absolute prices in the data set. Second, the decomposition of ΛF_t might not be unique. That is, a_t and R_t might not be identifiable separately due to difficulties in distinguishing between absolute and average relative price changes. Instead, Reis and Watson (2007) propose their pure inflation measure:[7]

$$v_t = a_t - E\left[a_t \mid \{R_\tau\}_{\tau=1}^{T}\right]. \tag{6}$$

The rationale for estimating pure inflation is that v_t is the common component in price changes that has an equiproportional effect across all prices and that it is not correlated with relative price changes. Expanding the interpretation, this measure of pure inflation turns out to be a proxy for core inflation since it isolates core price changes from relative shifts. Despite that pure inflation estimation is different to standard core inflation calculations (in which highly volatile CPI prices are excluded), its measure could be an alternative indicator of nominal inflation. However, the need to revise past pure inflation values whenever a new estimation is done prevents using the index as a standard core inflation measure.

4 Estimating the model

In order to estimate the model, we follow the three-step approach of Reis and Watson (2007). First, estimation of the number of factors k. Secondly, estimation of factors a_t and R_t and the factor loadings (matrix Γ). Also, unit restrictions on the loadings in the absolute-price component (β) are evaluated. Lastly, the measure v_t of pure inflation is calculated.[8] Here, on top of those stages, some robustness analysis is also conducted.

In selecting the size of the vector k, we follow Bai and Ng (2002). This method rests on the number of dominant eigenvalues from the covariance

[7] An additional assumption must be made here. Loadings connected with factors responsible for relative price changes sum up to zero for all goods. This assumption is equivalent to saying that $\sum_{i=1}^{N} \gamma_{ij} = 0$, for $j = 1, 2, ..., k-1$, where γ_{ij} denotes the element (i, j) of matrix Γ.

[8] Details on the procedure could be seen in Reis and Watson (2007). We are grateful to them for providing us with their estimation codes.

matrix from the data. Based on our dataset, two and three factors are used alternatively for robustness.

Parameter estimation for equations (4) and (5) is done through an unobserved components model with behavioral assumptions on the latent variables a_t, R_t and u_{it}. It is assumed that (a_t, R_t) follows a VAR process, that u_{it} is an independent AR process, and that all errors are normal.

Thus, the following model of unobserved components is estimated:

$$\pi_{it} = \beta_i' a_t + \gamma_i' R_t + u_{it}, \tag{7}$$

$$\Phi(L) \begin{pmatrix} a_t \\ R_t \end{pmatrix} = \epsilon_t, \tag{8}$$

$$\rho_i(L)\, u_{it} = \alpha_i + e_{it}, \tag{9}$$

where $\{e_{it}\} \sim N\left(0, \sigma_i^2\right)$, $\{e_{jt}\} \sim N\left(0, \sigma_j^2\right)$, $\{\epsilon_t\} \sim N(0, Q)$ are mutually and serially correlated. Columns in Γ are normalized to identify the factors (they are orthogonal to each other). The EM algorithm (expectations maximization) by maximum likelihood is used for estimation. Estimates of $\Phi(L)$, then, allow to calculate the expectations considered in equation (6) to obtain pure inflation, v_t.[9]

5 Empirical evidence in Peru

Our database contains monthly information on price indices of 55 goods categories in the CPI at 3-digit disaggregation level. Price changes are estimated as $1200 \times [\ln P_{it} - \ln P_{it-1}]$, so that all observations are taken as monthly percentage change. Data sample spans from 1995:1 to 2008:7 (163 observations). Some price categories were left out from the actual estimation database. Three of those categories were not considered due to a clear seasonal pattern. Other seven items were excluded because of

[9] Another approach to estimate (4)-(5) is to use a restricted principal components model. It consists of solving the restricted least-squared problem: $min_{(\Gamma,a,R)} \sum_{i=1}^{N}\sum_{t=1}^{T}(\pi_{it} - a_t - \gamma_i' R_t)^2$. When N and T are large and the errors terms u_{it} are weakly cross-sectionally and serially correlated, this approach gives consistent estimators and the sampling error in the estimated factors is sufficiently small that is can be ignored (i.e. $\widehat{a}_t$ and $\widehat{R}_t$ may be used in regressions in place of the true values of a_t and R_t); see Stock and Watson (2002), Bai (2003), Bai and Ng (2006). We also use this approach in this paper. The estimates are very similar to those obtained with the unobserved-components model but more volatile. Results are available upon request.

their excessive number of zero-change observation values (above 30 percent of observations). Thus, $N = 45$ are the prices considered for estimation. We searched statistically for outliers and in a few cases original values were replaced by local media.[10]

The model (4)-(6) imposes the restriction that loadings on the absolute-price factor must be one for all goods. To investigate how restrictive this is, we calculate the increase in fit that comes from dropping the restriction, measured as the fraction of (sample) variance of π_i explained by the factors. Moreover, we estimate the value of θ_i in the N regressions:

$$\pi_{it} = \theta_i a_t + \lambda_i' R_t + u_{it}, \tag{10}$$

using $\widehat{a}_t$ and $\widehat{R}_t$ instead of a_t and R_t. When $\theta_i = 1$, this corresponds to the restricted model, so we can use estimates of θ_i to judge how adequate this restriction is.

Figure 1 shows the largest twenty eigenvalues of the sample correlation matrix. It is clear that there is one large eigenvalue, but it is much less clear how many additional factors are needed. Estimates of Bai and Ng (2006) suggest one or two factors (see Table 1).[11] The results show that the factor connected to the first eigenvalue is responsible for around 23% of volatility of individual inflation rates. Other factors lead to an average of share in variance less than 5.5%.[12] In our benchmark model we use 2 factors: a_t and R_t.

Figure 2a shows the fit of unrestricted factor models that do not impose the unit restriction on the loading of the absolute-price factor. The unrestricted models appears to fit better only for a small number of price series.

Figure 2b shows the ordered values of the estimates of θ_i, that is the least-squares coefficient from regressing π_{it} on $\widehat{a}_t$ controlling for $\widehat{R}_t$. Most of the estimates are close to 1 which means that the restriction is potentially valid. Figure 2c shows the ordered values of the (4 lag Newey and West) t-statistic testing that $\theta_i = 1$. We obtain 11.1% of the t-statistics above the

[10] The model in(4)-(6) does not require any expenditures shares, since the objective of measuring pure inflation is not to measure the cost of living, but rather to separate absolute from relative price changes.

[11] The results of Reis and Watson (2007), using the information criteria of Bai and Ng (2006) suggest 2 factors, one factor, or 11 factors (according to the ICP_1, ICP_2, and ICP_3, respectively). In the case of Brzoza-Brzezina and Kotlowski (2009), using Polish data, the information criteria suggested 12 factors, 9 factors, and 12 factors, respectively. Despite these results, they conclude that the true number of factors should be equal to 2 or 3.

[12] First four factors account for around 39% of volatility of individual inflation rates.

standard 5% critical values and 8.9% above the 1% critical values. Therefore, the results suggest that the null hypothesis of $\theta_i = 1$ cannot be rejected, that is, little is lost by imposing this restriction.[13]

Robustness analysis suggests that VAR order (we tried using 4, 6, and 8 lags) does not change results importantly. Also, our results are robust to the stochastic behavior of a_t and R_t. Indeed, we use different combinations, but the preferred one is with $a_t \sim I(1)$ and $R_t \sim I(0)$.

Absolute price changes closely determine the inflation rate dynamics and relative price movements influence the inflation level[14]. Importantly, after the adoption of the inflation targeting regime in 2002, volatility in absolute and relative price changes has decreased (in particular, uncertainty in absolute price movements). The measure of pure inflation shows three different patterns over the sample. First, there is a decreasing inflation rate from 1995 until the end of 2001 that corresponds to a middle-inflation fighting monetary policy. At this stage, the central bank was targeting money aggregates. Second, it follows a period of low and stable inflation rate after monetary policy shifted to directly targeting inflation rates. Lastly, there is an upsurge in inflation rates to slightly above-the-target levels, from 2007 onwards, that might reflect both domestic demand pressures and international cost shocks.

Interestingly, estimates of pure inflation closely resemble behavior from core inflation, though pure inflation seems to be smoother. Figure 3a plots core inflation and a measure of pure inflation with our preferred scenario. It seems that when absolute price changes are allowed to be persistent, $a_t \sim I(1)$, and relative price changes are mean-reverting, $R_t \sim I(0)$, the measure of pure inflation is more closely related to the standard measure of core inflation. Indeed, the correlation between this measure of pure inflation and a standard core inflation is large and highly significant (0.989). In the other two cases presented here, pure inflation still follows core inflation dynamics but some discrepancies appear. Figure 3b assumes $a_t \sim I(1)$ and $R_t \sim I(1)$ and it provides a measure of pure inflation which is significantly correlated to core inflation (0.977). Figure 3c shows results with $a_t \sim I(0)$ and $R_t \sim I(0)$ and it also shows a highly significant correlation between pure and core inflation (0.957).[15]

[13] Reis and Watson (2007) find that over 30% of the t-statistics are above the standard 5% critical values and over 20% above the 1% critical value. Despite these large values, they accept the restriction.

[14] These results are not in contrast to Castillo, Humala, and Tuesta (2006), who document the non-linear link between inflation levels and the variance of both permanent and transitory components of inflation for an extended sample in Peru (including the hyperinflation episode).

[15] Pure inflation estimates are more volatile compared with estimates obtained using the

6 Conclusions

Data information at a fairly disaggregated level of price categories in the CPI can be reduced to a few factors with economic interpretation. Thus, inflation could be decomposed into an absolute change measure and a relative change part. Correspondingly, a measure of pure inflation could also be estimated as a representation of the equiproportional effect across all prices that is not correlated with relative price changes. Since this pure inflation measure isolates inflation from aggregate and idiosyncratic relative price changes, it provides a rationale for an alternative measure of core inflation. Indeed, in the case of Peru, pure inflation estimates are highly correlated to standard measures of core inflation, despite the fact that they are very differently calculated. Empirical results are robust to different specifications with respect to lags and to the stochastic behavior of the estimating factors.

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principal components approach. Nonetheless, results and conclusions are fairly similar and the results are available upon request.

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Table 1. Selection of the Number of Factors

Number of Factors	Eigenvalues	Difference of Eigenvalues	ICP_1	ICP_2	ICP_3
0			8.894	8.894	8.894
1	10.170	7.527	8.737	8.744	8.721
2	2.643	0.214	8.758	8.772	8.725
3	2.428	0.427	8.781	8.801	8.731
4	2.001	0.135	8.811	8.839	8.745
5	1.866	0.121	8.842	8.876	8.759
6	1.745	0.232	8.872	8.914	8.774
7	1.513	0.049	8.908	8.956	8.793
8	1.463	0.106	8.941	8.996	8.809
9	1.357	0.048	8.975	9.037	8.827
10	1.309	0.072	9.007	9.076	8.842
11	1.236	0.068	9.037	9.114	8.857
12	1.167	0.094	9.067	9.151	8.870
13	1.072	0.041	9.098	9.188	8.845
14	1.031	0.031	9.127	9.224	8.897
15	0.999	0.057	9.153	9.256	8.906
16	0.942	0.071	9.177	9.287	8.913
17	0.871	0.054	9.201	9.318	8.921
18	0.816	0.048	9.224	9.348	8.928
19	0.768	0.060	9.246	9.377	8.933
20	0.707	0.031	9.268	9.406	8.939

Second column features the 20 largest eigenvalues of the correlation matrix arranged in decreasing order. Third column denotes differences between them. The last three columns feature values of information criteria as suggested by Bai and Ng (2002).

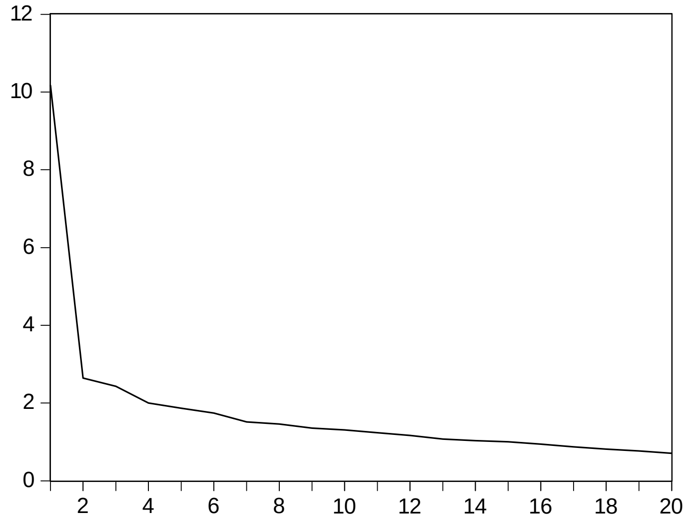

Figure 1. Eigenvalues of the Correlation Matrix

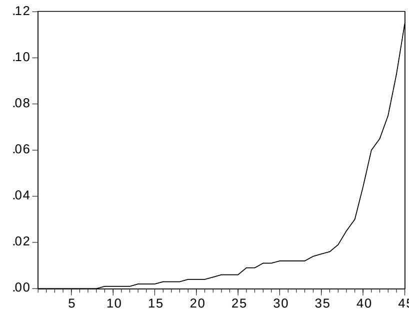

Figure 2a. Increase in R^2 from moving to Unrestricted Model

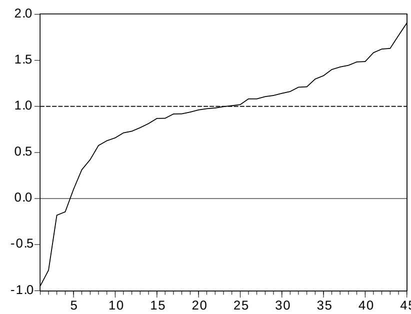

Figure 2b. Estimates of $\beta_i = 1$ (Coefficient on Absolute-Price Component)

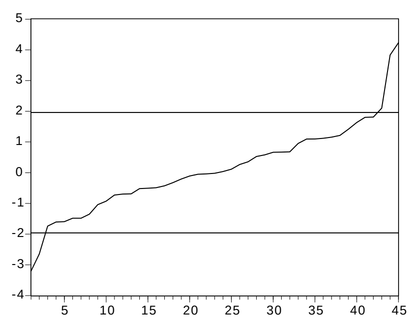

Figure 2c. Individual t-statistics for Hypothesis $\beta_i = 1$

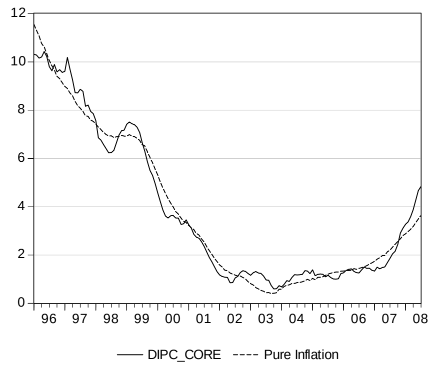

Figure 3a. $a_t \sim I(1)$, $R_t \sim I(0)$

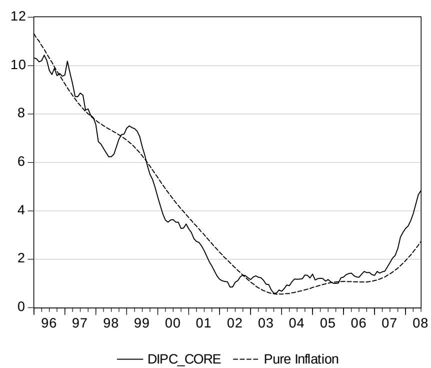

Figure 3b. $a_t \sim I(1)$, $R_t \sim (1)$

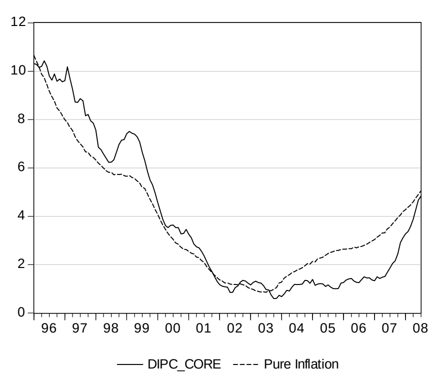

Figure 3c. $a_t \sim I(0)$, $R_t \sim I(0)$

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Av. Universitaria 1801, Lima 32 – Perú.
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