Core Inflation and Inflation Targeting in a Developing Economy

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Working Paper No. 02-W07

May 2002

DEPARTMENT OF ECONOMICS
VANDERBILT UNIVERSITY
NASHVILLE, TN 37235

www.vanderbilt.edu/econ

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*I am thankful to Mario Crucini, Huberto Ennis, Barry Eichengreen, James Foster, Ravi Kanbur, Karl Shell, Tim Vogelsang, and seminar participants at Vanderbilt University and the V meeting of the Network of America Central Banks Researchers for their comments and suggestions. I also thank José J. Rojas at Banco Central de Nicaragua for providing me with the data. The usual disclaimer applies. Support from the Center for Analytic Economics and the Thorne Fund at Cornell, and from Banco Central de Nicaragua is gratefully acknowledged. **Corresponding Address**: Department of Economics; Box 351819 Station B; Vanderbilt University; Nashville, TN 37235; **E-mail**: luis.rivas@vanderbilt.edu

ABSTRACT: This paper is concerned with inflation targeting as a potential monetary policy objective in a developing economy. Using data from Nicaragua, it first studies the extent to which the Consumer Price Index (CPI) could be used to formulate short–run inflation targets. It is found that due to the particular cross–sectional properties of the relative–price distributions, the rate of change in the CPI may not be the *best* index for this purpose. As a consequence, the paper is also concerned with the choice of alternative indicators of inflation and their statistical properties. These alternative measures are ranked according to their ability to forecast the rate of change in the price level. Finally, the relationship between the dispersion and skewness of the relative–price distribution and generalized inflation is studied using time series analysis.

Keywords: Core inflation, developing economies, forecasting, monetary policy, price index, time-series

JEL: C22, C43, E31, E37, E52, E58, O23, O54

1 Introduction

The principle that policy should be aimed at stabilizing some aggregate nominal variable enjoys considerable agreement in modern monetary economics. However, disagreement prevails over the variable to be stabilized. Some economists argue that the rate of change of some monetary aggregate should be chosen (Friedman, 1969). Others advocate the use of nominal income as the target for policy (Hall, 1985; Taylor, 1985). Yet, others maintain that the objective of monetary policy should seek to target inflation (Mishkin, 1999; Svensson, 1999; Eichengreen, 2001). This lack of consensus stems from the complexity of ranking and comparing the various trade–offs that the different regimes entail.[1]

A growing number of countries have shifted toward the maintenance of low inflation rates as their primary monetary policy objective.[2] Most of these countries have set specific target bands for the rate of inflation and have publicly committed to attain them. Besides providing a nominal anchor that ties down inflation expectations,

[1] There is even more agreement over the principle that, given the nominal variable, the stabilization policy should be implemented under a rule rather than unconstrained discretion. For theories stating that there could be substantial gains, mainly in terms of stability of output and prices, when the monetary authority credibly commits to a set of rules, see Kydland and Prescott, (1977) and Barro and Gordon (1983). For a comparative analysis of different target variables, see Mishkin (1999).

[2] Some of the countries that have explicitly targeted inflation are Australia, Canada, Finland, New Zealand, Sweden, Spain, and the United Kingdom. Chile and Israel have simultaneously targeted the exchange rate and the inflation rate. For a survey on the experience in targeting inflation, see Kahn and Parrish (1998).

advocates of inflation targeting maintain that this policy improves transparency and accountability because it conveys a precise, and readily understood goal. An important aspect of inflation targeting, which is of particular interest in this study, is that deviations from targets are routinely allowed in response to supply shocks: indexes of *core* or *underlying* inflation are often used to exclude or moderate the effects of these shocks (Mishkin, 1999). Opponents of inflation targeting question the desirability of this policy based on the difficult tasks of forecasting inflation and formulating mechanisms that could measure the feedback that such policy may have on inflation itself (Cecchetti, 1995; Woodford, 1994).

After the implementation of a stabilization plan in 1991 and the subsequent taming of high inflation rates, central bank officials in Nicaragua have become increasingly interested in further reducing the level of inflation and its variability.[3] This concern has not yet translated into concrete policy actions, but is reflected in a sustained effort for improving traditional measures and contemplating the introduction of measures of core inflation that could be potentially used as benchmark for policy purposes

[3] A successful stabilization program was implemented in March 1991, after a period of extremely high inflation. The course of events is as follows. There was an unsuccessful attempt to stabilize the economy in the late 1980s followed by a hyperinflationary period, attributed to fiscal imbalances caused by a civil war. The economy was anchored in 1990 with a new currency that was exchanged for the dollar at par. Towards the end of 1993, however, low international reserves caused an appreciation of the real exchange rate, which lead to a devaluation in January 1994, to the adoption of a crawling peg, and to a structural adjustment in April of the same year.

(Rivas and Rojas, 2001). The improvement of existing measures –like the consumer price index (CPI)– has been made possible through the collection of more reliable data since 1990. Availability of this data is the result of a renewed effort to increase the quality of national statistics and to a new awareness of the importance of central bank credibility and the transparency of monetary policy. The potential of inflation targeting as an alternative policy has thus shifted officials' attention toward the question of how to best measure core inflation.

Implicit in the above discussion is the idea that "headline" or CPI inflation is fundamentally different from the inflation pertinent for policy making. The point of departure for this view is the perception that individual prices have an underlying or core component, mostly associated with expectations and monetary expansion, and an idiosyncratic component. Regardless of whether the idiosyncratic component is transient or permanent in nature, one would expect a temporary impact on measured inflation, unless monetary policy validates the change in inflation, rather than just the change in the price level associated with the shock (see Rogers, 1998).

A common practice aimed at identifying core inflation is to remove certain categories of prices from those used in the construction of the CPI. The prices of food and energy, labelled as "too volatile", are suppressed while the remaining groups are reweighted to construct a measure of core inflation alongside the CPI. Another method seeks to exclude the effects of changes in indirect taxes, on the basis that indirect taxes ought not to be of concern for monetary policy purposes (Wynne, 1999).[4] In brief, individual prices are reweighted in terms of the strength or quality of their inflation signal, rather than by their expenditure shares. Other approaches calculate robust measures of inflation. These measures also suppress some items used to calculate the CPI. It is argued, however, that less information is lost in the process and that the criteria under which such items are chosen is statistically sound (Bryan and Cecchetti, 1994). Still other methods perceive core inflation as the component of measured inflation that has no medium to long term impact on real output. This approach uses other aggregate variables in the construction of measures of core inflation (Quah and Vahey, 1995).[5]

This study is primarily concerned with two aspects associated with inflation targeting as a potential monetary policy objective in developing countries like Nicaragua: the construction of measures of core inflation, and the evaluation of their forecast ability. What warrants the study of indexes of core inflation and their usefulness in the formulation of monetary policy are at least two issues. The first one is empirical: underlying inflation measures have been widely used in countries that have targeted inflation explicitly.[6] The second is statistical: as will be discussed at length below, various empirical studies document that relative prices are often not normally distrib-

[4] For methodological issues, see Donkesrs and others (1983) and Diewert and Bossons (1987).

[5] Rogers (1998) and Wynne (1999) review the concept and measurement of core inflation.

[6] New Zealand, Canada, United Kingdom, Finland, and Australia use or have used measures of underlying inflation for stipulating their inflation targets, at least in the short run. See Kahn and Parrish (1998).

uted. In this case the CPI may not be the most efficient estimator of core inflation, and other indexes need to be explored.

The non-normality of price relatives –the evidence that inflation is significantly correlated with the variance and skewness of individual price changes– has motivated important theoretical studies relating relative price changes and aggregate inflation. Two prominent papers are Ball and Mankiw (1995) and Balk and Wynne (2000). Ball and Mankiw put forth a menu-cost model to show that when a shock affects its relative price, a firm will adjust its price only if the desired adjustment is sufficiently large as to warrant paying the adjustment cost. Since firms respond more quickly to large shocks than to small ones, the desired increases occur more quickly than the desired decreases. According to the authors, that explains the positive correlation of changes in the price level with the standard deviation and skewness of the relative price distribution. Balk and Wynne, on the other hand, contend that little evidence exists to support the argument that such relationships arise from price rigidities. They argue that technology shocks themselves could be responsible for the positive inflation-skewness correlation. They provide a model that explains that such a relationship occurs because the same correlation is also present in the underlying sectoral technology shocks.[7]

The theoretical explanations above are based on empirical evidence arising from

[7] There are other competing explanations: the asymmetrical price response hypothesis, and the Lucas-type information imperfections theory. See Balk and Wynne (2000) for a more complete review of this literature.

OLS regressions of inflation on the standard deviation, the skewness of relative prices, and on other variables that proxy the shape of such distributions. Potential non-stationarity of some of these variables raises the possibility that such regressions may lead to incorrect inference. Thus, another concern of this paper is to study the Nicaraguan time series using appropriate econometric techniques.

The organization of the paper is as follows. Section 2 describes the data set and introduces some basic notation. Section 3 explain why it is appropriate to use measures of core inflation to study inflation from the monetary policy standpoint, and explores some salient features of the data. In Section 4, various measures of core inflation are computed. Section 5 presents the time series analysis. Section 6 ranks the various measures of core inflation. Comments and conclusions are offered in Section 7.

2 Data source and notation

2.1 The data

The available data set is a balanced panel composed of 27 *groups* of prices used for computing the CPI for the period of January 1988 to December 1998. Each price *group* constitutes an agglomeration of prices of different products that share a common characteristic.[8] For example, group one corresponds to cereals and its

[8] Between 1999 and 2000, the aggregation classification changed when a new CPI was introduced. However, the historical series were not recomputed with the new classification; therefore the old

products, and is composed of three *items*: (*i*) corn and its products, (*ii*) wheat and its products, and (*iii*) rice. Each of these items, in turn, is an aggregation of several *goods*. It is worth mentioning that the various groups are aggregated into *chapters*, and the aggregation of chapters results in the general price level. As an example, the first chapter is composed of the following groups: (*i*) cereal and its products; (*ii*) meat, fish and poultry products; (*iii*) milk, lacteal products, and eggs; (*iv*) oils and fats; (*v*) Fruits and vegetables; (*vi*) Sugars and sweets; (*vii*) Non–classified food; (*viii*) beverages and its products; and (*ix*) food away from home. In brief, in Nicaragua the categories employed are, from the highest level of disaggregation to the highest level of aggregation, goods, items, groups, chapters, and finally the CPI. The data were obtained from the national statistics office and the central bank of Nicaragua.

Figures 1 and 2 plot the month to month inflation and twelve-month inflation rates, respectively. That is, Figure 1 plots the inflation that took place between two consecutive months, say between 1991:11 and 1991:12. Figure 2, on the other hand, graphs the inflation rate between the same month of two consecutive years, say between 1990:12 and 1991:12. Notice that inflation was quite high and volatile during the years 1988-1991. This period is followed by substantially lower and less volatile inflation. Also, note that twelve-month inflation was about 400% in January 1989, after a hyperinflationary period, but it was declining as a result of an attempt to stabilize prices, undertaken in 1988 (Figure 2). Inflation had decreased to 200%

classification was kept.

by February 1990, only to increase to about 500% in February 1991, and then fell abruptly with the new stabilization program of March 1991. Since 1992, twelve-month inflation has remained below 20%.

It is clear from Figure 1 that the data generating processes for inflation in the CPI differs substantially before and after the stabilization plan of March 1991. Thus, the focus will be on the post-stabilization period (1991:06 to 1998:12). Another important reason to perform the analysis with the post-stabilization period is that until the end of 1990, most prices were, in one way or another, controlled by the government, and thus insulated from a variety of economy-wide as well as sector-specific shocks.[9]

2.2 Notation

In what follows, p_{it} will denote the headline or CPI price index of group i at date t, and P_t represents the aggregate price level or CPI index value at date t. Over a horizon τ, inflation in an individual group price is given by

$$\pi_{it}^{\tau} = \frac{1}{\tau}\left(\frac{p_{it} - p_{it-\tau}}{p_{it-\tau}}\right) \tag{1}$$

Similarly, over a horizon τ, mean inflation per period is given by

$$\Pi_t^{\tau} = \sum_i r_{it}\pi_{it}^{\tau} \tag{2}$$

where r_{it} is the relative importance weight of the i^{th} group at date t. Whenever τ is omitted, the horizon is taken to be one month ($\tau = 1$). In Nicaragua, the CPI

[9] Some prices remain controlled, like gasoline, collective urban and inter–urban transportation, and communications.

is a Laspeyres–type index with fixed weights, which for group i is denoted by w_i, with the property, of course, that $\sum_i w_i = 1$. It is easy to show that in such a case $r_{it} = \frac{w_i p_{it-\tau}}{\sum_i w_i p_{it-\tau}}$.[10] For future reference, note that $\sum_i r_{it} = 1$. Additional notation will be introduced as needed.

3 Consumer prices and core inflation

The point of departure in the construction of all measures of core inflation is that individual price changes share a common component that constitute core inflation, and an idiosyncratic component that primarily reflects sector specific developments. The task then is to isolate these two components of observed price changes. Using the notation developed so far, one can write this formally as

$$\pi_{it} = \pi_t^c + z_{it} \tag{3}$$

where π_t^c stands for the common or core component and z_{it} stand for the idiosyncratic component of the price change of group i at date t. Note that under the assumptions that the idiosyncratic component is normally distributed with $E(\mathbf{z}_t) = 0$ and $E(\mathbf{z}_t\mathbf{z}_t') = \sigma_t^2\mathbf{I}_n$, (where $\mathbf{I}_n$ is the $n \times n$ identity matrix and n the number of price

[10] To see this note that if the aggregate price level at time t is given by $P_t = \sum_i w_i p_{it}$, average inflation per period over a horizon τ is given by $\frac{1}{\tau}\left(\frac{P_t - P_{t-\tau}}{P_{t-\tau}}\right) = \frac{1}{\tau}\left(\frac{P_t}{P_{t-\tau}} - 1\right)$. Take now $\frac{P_t}{P_{t-\tau}}$ and notice that

$$\frac{P_t}{P_{t-\tau}} = \frac{\sum_i w_i p_{it}}{\sum_i w_i p_{it-\tau}} = \frac{\sum_i w_i p_{it-\tau}\,(p_{it}/p_{it-\tau})}{\sum_i w_i p_{it-\tau}} = \sum_i r_{it}\frac{p_{it}}{p_{it-\tau}}$$

and clearly, $\frac{1}{\tau}\left(\frac{P_t}{P_{t-\tau}} - 1\right) = \frac{1}{\tau}\left(\sum_i r_{it}\frac{p_{it}}{p_{it-\tau}} - 1\right) = \Pi_t^\tau$, which follows because $\sum_i w_i = 1$.

groups) the maximum likelihood estimator (MLE) of core inflation at date t is the (arithmetic) mean inflation, $n^{-1}\sum_i \pi_{it}$.[11]

However, as Wynne (1999) points out, it is not necessarily the case that all prices are equally informative about inflation and thus equally important. Diewert (1995) shows that for the weighted average $\sum_i r_{it}\pi_{it}$ to be a MLE of core inflation, the variance assumption must be replaced by $E(\mathbf{z}_t\mathbf{z}_t') = \sigma_t^2\mathbf{R}_t^{-1}$, where $\mathbf{R}_t = diag\,[r_{1t}, ..., r_{nt}]$; that is, the variance of each idiosyncratic component must be inversely related to its corresponding relative importance, provided that the latter is fixed and non-random for each individual inflation.

Multiplying (3) by r_{it}, and aggregating across groups, one obtains

$$\sum_i r_{it}\pi_{it} = \pi_t^c + \sum_i r_{it}z_{it} \tag{4}$$

Thus, if the term $E\left(\sum_i r_{it}z_{it}\right)$ is zero, which will be the case under the assumptions given above, then inflation given by the CPI would be the MLE of core inflation. However, studies for various countries show that such term is generally nonzero. In fact, the cross-section distribution of price changes are typically not normally distributed. Rather, individual prices are characterized by leptokurtic and skewed distributions, at least for high frequency inflation data.[12] A natural next step will be

[11] In the stochastic approach to price indexes, this index is known as the Carli (1764) index. For the properties of the Carli index, see Diewert (1995)

[12] See Bryan and Cecchetti (1994), Bryan, Cecchetti and Wiggins (1997), Bryan and Cecchetti (1998), and Vining and Elwertowski (1976) for the US; Rogers (1998) for New Zealand; Mio and Higo (1999) for Japan; Lach and Tsiddon (1992) for Israel; among others.

to attempt to quantify $\sum_i r_{it} z_{it}$ for the Nicaraguan price data.[13]

Figure 3 plots aggregate month to month percentage change in the price level, $\frac{P_t - P_{t-1}}{P_{t-1}} \times 100$ (left scale), along with the dispersion of individual price changes for the period in question, defined by $SD_t = \left(Var(\frac{p_{it} - p_{it-1}}{p_{it-1}}) \right)^{1/2}$ (right scale). It can be observed that with the implementation of a structural adjustment program in 1994, $\{SD_t\}$ seems to have experienced a moderate decline, which lasts for most of the remaining sample period. It is more difficult to discover a consistent relationship between the dispersion of individual price changes and CPI inflation, though their correlation over the period is around 0.45.

Figure 4, which plots inflation (left scale) and the skewness of the cross-sectional distribution of price changes, denoted by S_t (right scale), suggests that the direction of skew is the same as the direction of change in the rate of inflation. For instance, if the aggregate rate of inflation is increasing, then the distribution seems to be positively skewed, with most groups' price changes below the rate of inflation and a few above but generally at a great distance from it. The opposite occurring when inflation declines. The correlation between these two variables is approximately 0.70.

[13] It should be mentioned that the term $\sum_i r_{it} z_{it}$ could be decomposed into bias and noise. The bias comes from the CPI's inability to fully account for substitution among goods, quality changes, and the introduction of new products. The analysis herein concentrates on the noise component assuming away measurement problems. An excellent theoretical treatment of the bias in the cases of consumer indexes and production indexes is offered by Fisher and Shell (1972) and Fisher and Shell (1998), respectively.

A well-established literature have found similar patterns in US data, and more recent research suggests that such patterns are present in other finite sample data.[14]

In an effort to better understand the properties of the cross–section distribution of price changes, the moments of the distributions at different horizons are studied. Following Bryan and Cecchetti (1999), define the n^{th}–order central (weighted) moment as

$$m_{nt}^{\tau} = \sum_{i} r_{it} \left(\pi_{it}^{\tau} - \Pi_{t}^{\tau}\right)^{n} \tag{5}$$

(i.e., $n = 2$ is the second-order moment or weighted variance).

Utilizing the price level data, panels of inflation rates at different horizons for the 27 price groups over the period of June 1991 to December 1998 are constructed. For each of these panels, the moments of the distribution are computed across groups at each date. Thus, for each panel, one obtains a time series for each of the distribution moments. Table 1 reports descriptive statistics for the moments of price changes at overlapping horizons of 1 to 36 months. In computing the moments, the actual expenditure weights are used. At each horizon τ, the table reports summary statistics for the weighted standard deviation, given by

$$WSD_{t}^{\tau} = \left(m_{2t}^{\tau}\right)^{\frac{1}{2}}, \tag{6}$$

[14] Vining and Elwertowski (1976) found similar patterns in annual US data in the period 1948-1974. For a review of the literature see Marquez and Vining (1984), and more recently Lach and Tsiddon (1992). See previous footnote for evidence supporting such hypothesis in countries like New Zealand, Japan, and Israel.

and the weighted skewness and kurtosis, which are the "scaled" third and fourth moments, given respectively by

$$WS_t^\tau = \frac{m_{3t}^\tau}{(m_{2t}^\tau)^{\frac{3}{2}}}, \quad WK_t^\tau = \frac{m_{4t}^\tau}{(m_{2t}^\tau)^2} \tag{7}$$

The most important feature of Table 1 is that the distributions of individual price changes are leptokurtic. At the monthly frequency, these distributions have, on average, a kurtosis larger than 6 and a large standard deviation, approximately 6.4. In addition, the kurtosis declines as the horizon over which it is averaged increases.

A second feature of these distributions is that, although they seem to be skewed at high frequencies, the skewness tends to decline as the horizon increases, implying that in the very long-run distributions of individual price changes tend to be less asymmetrical. For example, at $\tau = 36$, mean skewness is merely 0.06 with a standard deviation of about 0.8. The summary statistics of Table 1 show that in the case of Nicaragua there is a potentially important source of high frequency noise in the measurement of inflation.

4 Measures of core inflation

The analysis in the previous section shows that the normality assumption is not borne out by the data. In this section, various measures of core inflation are constructed. The measures considered belong to two major classes: (i) measures that weight the various price groups based on their inflation signal, and (ii) measures that are robust to departures from normality. These are not the only measures of core inflation found in the literature. There are three main reasons for restricting the analysis to these two classes of underlying inflation measures.

First, such measures should be computable in real time. This criterion leaves out measures based on filtering methods. Even when some filtering techniques are used to compute real time measures, like in the well-known case of the Hodrick–Prescott filter, the end of sample adjustments make them unappealing (Baxter and King, 1995).

A second criterion is that the measure of core inflation be "history independent". Measures whose past values need to be recomputed every time new observations obtain are not considered, primarily on the basis that it becomes hard for the central authority to justify to the public such changes, especially when inflation targeting is pursued under a rule. This condition leaves out measures like the dynamic factor index model proposed by Bryan and Cecchetti (1993), which uses contemporaneous as well as previous information of individual price changes.[15]

The final criterion used for choosing the inflation measures analyzed below is that only price data be used for their construction. This condition excludes the approach proposed by Quah and Vahey (1995), who define core inflation as the component of measured inflation that has no impact on real output in the long–run. This measure is constructed by imposing long–run restrictions on a bivariate VAR system of output

[15] This index brings to bear the persistence of both individual price changes and general inflation. However, it makes strong restrictive assumption on the idiosyncratic component of price disturbances at all leads and lags.

and inflation, and is motivated on the notion of a vertical long–run Phillips curve.[16]

4.1 The quality of the inflation signal

Many statistical agencies use measures that exclude certain items such as food, energy, and indirect taxes, under the criterion that they provide little or no information about underlying inflation. The idea is that items should be weighted in accordance to the strength or quality of their inflation *signal*.

One of the measures evaluated is the one used by central bank officials in Nicaragua as a measure of core inflation. This measure removes the price groups that correspond to food and energy, and will be denoted by I_X.[17] It is constructed as follows. The food and energy items are removed from the individual price change data given by the CPI, by setting their corresponding weights to zero. The remaining groups are reweighted. That is, if groups k and s are removed, for each remaining group with initial weight w_i, the new weight used to construct I_X is $\frac{w_i}{\sum_{j \neq k,s} w_j}$.[18] Figure 5 presents

[16] The first and second criteria imposed on the set of underlying inflation measures above seem reasonable from the policy standpoint. The third is more questionable, especially because the Quah–Vahey measure adopts a theoretical framework in its construction and is not solely based on statistical grounds. Unfortunately, in the case of Nicaragua, output data are either unreliable during some of the period or unavailable, making it hard to relax this criterion.

[17] More precisely, it removes all food items, energy (electricity), potable water services, and communications (telephone services). Water, telephone services, and electricity are removed under the criterion that they are produced by public monopolies whose prices are administered. The main idea is that they seldom respond to demand pressures.

[18] It could be argued that removing these price groups is a questionable procedure because together

I_X along with CPI inflation.

The second measure studied is the one proposed by Dow (1994) and Diewert (1995), which is implemented by Wynne (1997) for US data, and is known as the New–Edgeworthian measure. Formally the measure is constructed as the solution to the following nonlinear system of equations:

$$I_{V,t} = \frac{\sum_i \frac{\pi_{it}}{\sigma_i^2}}{\sum_i \frac{1}{\sigma_i^2}}, \quad \sigma_i^2 = \frac{1}{T} \sum_t (\pi_{it} - I_{V,t})^2 \tag{8}$$

where $I_{V,t}$ and σ_i^2 are the unknowns (for each t and i, respectively, where $t = 1, ..., T$). This measure will be denoted by I_V. Notice that in this case the weights are chosen so that individual price changes are inversely proportional to the volatility of those prices. It is claimed that its main advantage over measures that arbitrarily exclude some price groups (i. e., food, energy, etc.) is that it retains information contained in the discarded price groups.[19] For the Nicaraguan data, the inflation measure I_V, along with rate of change in the CPI are plotted in Figure 6.

food and energy represent an unusually large fraction of the expenditures of the average household, as reflected in the expenditure weights utilized to aggregate price groups in the calculation of the CPI. At the first level of disaggregation, the category of food and beverages is assigned a weight of 0.46. If we add gas and electricity (not including gasoline), they constitute about 59% of total expenditures.

[19] If the variance of the idiosyncratic component of each price group is assumed constant over time, I_V is equivalent to a measure resulting from the estimation of π_t^c using a generalized least squares (GLS) procedure.

4.2 Robust Estimators: Weighted α–trimmed means

A criticism to the removal of food and energy is that it may be the case that the frequency at which food and energy items lie in the tail of the distribution, therefore having poor quality in terms of their inflation signal, may be surprisingly low.[20] This type of criticism, along with the observed non-normality of the individual price change distribution have given rise to measures of inflation that are robust to departures from normality and which give a statistical criterion (but not economic) to the removal of the items with poor quality in their inflation signal (Bryan and Cecchetti, 1993, 1994). The basic idea is that, given the characteristics of the distribution of individual prices explained in the previous section, these are more robust measures of central tendency than the CPI. The α–trimmed estimators belong to the class of linear combinations of order statistics studied by Huber (1981) and others. Bryan, Cecchetti, and Wiggins (1997) show how these measures are constructed for price index data. To calculate the α–trimmed mean of individual price changes, individual inflation rates $\{\pi_{1t}, ..., \pi_{nt}\}$ are ordered along with their respective weights $\{w_1, ..., w_n\}$.[21] Define $W_i \equiv \sum_{j=1}^{i} w_j$.

[20] For the United States, Bryan, Cecchetti and Wiggins (1997) find that the frequency at which food away from home, one of the subcategories of food, lies in the tail of the distribution is quite small. For example, when they truncate the distribution of price changes, trimming 9% of the tails, food away form home is suppressed only about 3% of the time, indicating that the quality of its inflation signal may be considerable good.

[21] Expenditure weights are used in the calculation of the trimmed means.

Consider now $\alpha \in [0, 100)$. Let T_α be given by

$$T_\alpha = \left\{ W_i \mid \frac{\alpha}{100} < W_i < 1 - \frac{\alpha}{100} \right\}$$

The weighted α–trimmed mean is given by:

$$I_\alpha = \frac{1}{1 - 2\frac{\alpha}{100}} \sum_{i \in T_\alpha} w_i \pi_{it} \tag{9}$$

Notice that when $\alpha = 0$, inflation in the CPI obtains; when $\alpha = 50$, the sample (weighted) median obtains. Intuitively, $\frac{\alpha}{100}n$ observations are removed from each end of the sample and the weighted mean of the rest is taken. But as Huber (1981) explains, the trimmed mean estimators do not throw away all of the information sitting in the discarded observations.[22] Bryan Cecchetti, and Wiggins (1997) present evidence that, among the class of trimmed means, the CPI is not the most efficient estimator of core inflation. Three weighted trimmed mean estimator are constructed: I_{15}, I_{25}, and I_{50}. The somewhat arbitrary choice of the α values was made based on the extensive Monte Carlo experiments of Hogg (1967). As an example, Figure 7 plots the 15% trimmed mean along with CPI inflation.

It is clear from a simple examination of Figures 5 to 7 that the various measures of core inflation differ considerably during important periods. Apparently, in months

[22] This is a technical issue. The details can be found in Huber (1981). The idea is that the α–trimmed means do what another type of robust statistic does: it does not throw away any of the outliers but just makes a correction. This statistic is the α–Winsorized mean, which corresponds to replacing the $\frac{\alpha}{100}n$ leftmost observation by the $\frac{\alpha}{100}n + 1$ observation, and the $\frac{\alpha}{100}n$ rightmost observation by the $n - \frac{\alpha}{100}n$ observation, and take the mean of this modified sample.

with high inflation, the differences among the various measures seem to be exacerbated. Before leaving this section, however, it will be useful to provide summary statistics for the various measures of core inflation. Table 2 presents the correlation matrix and summary statistics for all the measures. On the basis of their correlation with CPI inflation, one can rank the measures as follows: I_V (0.88), I_{15} (0.68), I_{25} (0.66), I_{50} (0.61) and I_X (0.60). In addition, notice that I_X and I_V indicate an average monthly inflation less but fairly close to CPI inflation (about 16% and 19% lower, respectively). Furthermore, inflation given by I_{15} is on average about 24% lower than CPI inflation, while I_{50} and I_{25} offer monthly inflation rates that are roughly 51% and 63% lower than CPI inflation. Similar patterns can be found with respect to the volatility of the various measures of core inflation. According to their standard deviations, the most volatile is I_X (about 7% lower than CPI inflation), while the least volatile is I_{25} (about 74% less than inflation in the CPI).

5 Time series analysis

A first step in the study of core inflation is to determine whether the various time series under study, headline (or CPI) inflation, individual price dispersion and skewness, and the various core inflation measures are stationary processes. A natural second step is to check whether such time series are changing over time gradually. A priori, there are reasons to believe that some of these processes may display structural changes in their trends: the data used belong to a transitional economy, that underwent major reforms in the monetary, fiscal, and external sectors over the period under study.

This section attempts to shed some light on the issues of potential nonstationarity and structural breaks. The methodology used is that of Vogelsang (1997). His methodology is useful for a number of reasons. First, it handles the presence of serial correlation in the errors. In addition, it uses the Augmented Dickey–Fuller (ADF) factorization for testing structural breaks, which makes his methodology relatively easy to implement empirically.

Suppose that the data generating process of a generic time series $\{y_t\}$, with a break in the trend at unknown time T_b^c, be given by the following structural model

$$y_t = \theta_0 + \theta_1 t + \varphi_0 DU_t + \varphi_1 DU_t \left(t - T_b^c\right) + v_t \tag{10}$$

$$A(L)v_t = e_t \tag{11}$$

where DU_t is an indicator function that takes the value of 0 for t less than or equal T_b^c and 1 otherwise, $A(L) = 1 - a_1 L - a_2 L^2 - ... - a_{k+1} L^{k+1}$, and L is the lag operator. The autoregressive polynomial $A(\cdot)$ is assumed to have at most one real valued root inside the unit circle with all others strictly outside the unit circle, and e_t is a white noise process, with zero mean and variance σ_e^2. What the model says is that if the structural break occurs at date T_b^c, the intercept and deterministic trend parameters are given by θ_0 and θ_1, respectively (up to date T_b^c inclusive). From date $T_b^c + 1$ onwards, as a result of the structural change, such parameters are given by $\theta_0 + \varphi_0$ and $\theta_1 + \varphi_1$, respectively. When the series under study does not contain a deterministic trend component, (10) is given by $y_t = \theta_0 + \varphi_0 DU_t + v_t$. Notice that the autoregressive process need not be stationary. In fact, whether the process $\{y_t\}$ follows a stationary or unit root process will be one of the interest in this section. For this, it will be convenient to factor the polynomial $A(L)$ according to the augmented Dickey-Fuller (ADF) procedure as $A(L) = (1 - \alpha L) - C(L)(1 - L)$, where $C(L) = \sum_{i=1}^{k} c_i L^i$, $c_i = -\sum_{j=i+1}^{k+1} a_j$, and $\alpha = A(1) = \sum_{j=1}^{k+1} a_j$. Applying this factorization to v_t and defining $\gamma = \alpha - 1$ gives

$$\Delta v_t = \gamma v_{t-1} + \sum_{i=1}^{k} c_i \Delta v_{t-i} + e_t \tag{12}$$

which can be used to rewrite (10) as

$$\begin{aligned}\Delta y_t &= \beta_0 + \beta_1 t + \delta_0 DU_t + \delta_1 DU_t \left(t - T_b^c\right) + \sum_{j=0}^{k} \eta_j \mathbf{1}\left(t = T_b^c + j + 1\right) \\ &\quad + \gamma y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + e_t\end{aligned}$$

where $\beta_0 = (1 - \alpha)\,\theta_0 + (\alpha - C(1))\,\theta_1$, $\beta_1 = (1 - \alpha)\,\theta_1$, $\mathbf{1}(\cdot)$ is an indicator function that takes the value of 1 at date $T_b^c + j + 1$ and 0 otherwise, and the parameters $(\delta_0, \delta_1, \eta_0, \eta_1, ..., \eta_k)$ are given by solving:

$$\begin{aligned}&\sum_{j=0}^{k} \eta_j \mathbf{1}\left(t = T_b^c + j + 1\right) \\ &= \mu \left[(1 - \alpha)\,\varphi_0 + \left(\alpha - \sum_{i=j+1}^{k} c_i \right) \varphi_1 - \delta_0 + \left[(1 - \alpha)\,\varphi_1 - \delta_1\right]\left(t - T_b^c\right) \right]\end{aligned}$$

The one–time dummy variable term, $\sum_{j=0}^{k} \eta_j \mathbf{1}\left(t = T_b^c + j + 1\right)$, is asymptotically negligible (Vogelsang, 1997); therefore, it is convenient to drop it from the model and consider

$$\Delta y_t = \beta_0 + \beta_1 t + \delta_0 DU_t + \delta_1 DU_t \left(t - T_b^c\right) + \gamma y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + e_t \tag{13}$$

in which case $\delta_0 = (1-\alpha)\varphi_0 + \left(\alpha - \sum_{i=j+1}^{k} c_i\right)\varphi_1$, $\delta_1 = (1-\alpha)\varphi_1$.

For reasons that will be apparent later, it is convenient at this time to take a step back. Suppose that one knows that each of the sequences studied do not have an unstable trend function. If this is the case, then $(\varphi_0, \varphi_1) = (0, 0)$, and the process $\{y_t\}$ can now be described by

$$\begin{aligned} y_t &= \theta_0 + \theta_1 t + v_t \\ A(L)v_t &= e_t \end{aligned}$$

where the first equation is replaced by $y_t = \theta_0 + v_t$ if the process does not have a deterministic trend component. In this case, applying the ADF factorization and using the notation introduced above, one eventually arrives at

$$\Delta y_t = \beta_0 + \beta_1 t + \gamma y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + e_t \tag{14}$$

and to $\Delta y_t = \beta_0 + \gamma y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + e_t$ if the time series does not have a deterministic trend component. Of course, this is the classical ADF regression, which allows one to test for the presence of a unit root. The regression (14) is estimated and the results are reported in Table 3. The null hypothesis is that $\gamma = 0$. The lag length k was chosen using the approach recommended by Perron and Vogelsang (1992). One first estimates (14) by using a maximal value of 10 for k. One then tests the significance of the coefficient on the last included lag by using a 5% two–tailed t–test. Asymptotic normality of the t–test is used to carry out inference. Asymptotic normality holds whether the errors are stationary or have a unit root. If this coefficient is significant the procedure is stopped. Otherwise, k is reduced by 1 and (14) is estimated again using $k = 9$. This continues until one either finds significance or until $k = 0$. When $k = 10$ and the coefficient on the 10th lag was significant, the maximal value is increased to 15 and the process is repeated. Asymptotically, this procedure yields the same distributions under the null hypothesis as when the order of the lag is known (see Perron, 1990).

Note that a time trend coefficient is included in the case of the log of the CPI only. Simple inspection of Figures 1-6 indicates that for $\{SD_t\}$, $\{S_t\}$, and the various measures of core inflation, the ADF equation (14) should be estimated without a time trend. Table 3 indicates that there is not enough evidence to reject the hypothesis of nonstationarity of the price level (given by the log of the CPI), but the results also show that the log of the price level is first difference stationary, indicating that inflation is a stationary process. It is also found that the unit root hypothesis cannot be rejected for the skewness of the distribution of individual prices, and for the core inflation measure I_X (though the latter is first–difference stationary [not shown in the table]). The rest of the time series reject the hypothesis of nonstationarity at the 1% significance level. These results are surprising for a number of reasons. The possibility of persistence in the skewness of the distribution of individual price changes suggests that a long-run relationship between overall inflation and the skewness of such distributions may be lacking. This raises the possibility of mispecifications in OLS inflation-skewness regressions, and suggests a certain degree of caution regarding theories based on evidence obtained from such regressions. In addition, the above result opens the possibility that inflation in the CPI and core inflation, as given by the measure excluding food and energy, I_X, lack a long run relationship, and questions the use of such measure as a potential indicator to set intermediate inflation targets in Nicaragua.

Since it has been shown (Perron, 1990) that when a shift in the mean is mispecified, the estimate of γ in (14) will be biased toward zero. In such a case, the fitted model cannot be distinguished from a model with no unit root and a shift in the mean. Thus, the next step is to test for structural breaks in the trend function, utilizing the procedure proposed by Vogelsang (1997). This amounts to estimating the more general model given by (10) and (11) (or $y_t = \theta_0 + \varphi_0 DU_t + v_t$, if one is certain that the time series in question does not have a deterministic time trend component), which corresponds to estimate (13). Since in reality, the exact break date is not known to the researcher, this procedure involves computing Wald statistics for a break in the trend over a range of possible break dates and taking the supremum and exponential averages of the statistics.

Under the null hypothesis of no structural change, $\delta = 0$ (follows from $\varphi = 0$), equation (13) is estimated using ordinary least squares (OLS) using a hypothetical break $T_b = [\lambda T]$, where T is the sample size and $\lambda \in [\lambda^*, 1-\lambda^*] \subset (0,1)$.[23] Let $W_T^p(\lambda)$ denote the Wald statistic for testing $\delta = 0$. Define the possible break dates $\Lambda = \{T_b^*, T_b^* + 1, ..., T - T_b^*\}$, where $T_b^* = [\lambda^* T]$. The parameter λ^* is called the amount of trimming. Vogelsang (1997) proposes three test statistics, two of which belong to the class of statistics proposed by Andrews and Ploberger (1994), given by,[24]

$$\begin{aligned} Mean\ W_T^p &= T^{-1} \sum_{T_b \in \Lambda} W_T^p(T_b/T) \\ Exp\ W_T^p &= \log \left[T^{-1} \sum_{T_b \in \Lambda} \exp\left(\frac{1}{2} W_T^p(T_b/T) \right) \right] \end{aligned}$$

and the third, proposed originally by Quandt (1960) and generalized by Andrews (1993), given by

$$Sup\ W_T^p = \sup_{T_b \in \Lambda}\ W_T^p(T_b/T)$$

Notice that, as Vogelsang (1997) explains, the *Mean* and the *Exp* statistics are optimal in the case in which the deterministic trend component is not included and $\{v_t\}$ is $I(0)$, but not when $\{v_t\}$ is $I(1)$ and/or the deterministic component is included. The *Sup* statistic does not belong to the class of optimal statistics; however, it is useful because it provides an estimate of the true break date ratio λ_c.

It should be mentioned that, as Vogelsang (1997) proves, the limiting distributions of the statistics presented above are nonstandard, thus the critical values used are

[23] Note that the hypothetical break date may differ from the true break date T_b^c.

[24] Notice that the parameter λ_c, associated with the true break date, is present only under the alternative hypothesis.

taken from Vogelsang (1997). One advantage of using the above statistics is that for the case in which the errors are $I(0)$ and purely $I(1)$, the limiting distributions are free of nuisance parameters and depend only on whether the deterministic trend component is included or not, and on the amount of trimming, λ^*. First, for each series, one must determine the amount of trimming, including the possible break date(s). Then, (13) is tested with a dummy variable that takes the value of zero up to the hypothetical break date. This is repeated for each date belonging to the potential break dates. Thus, a series of Wald statistics is obtained for each of the original series. Then, these three statistics are constructed and their significance determined according to critical values tabulated in Vogelsang. The lag–length k for each regression is chosen according to Perron (1990), as explained above. Table 4 presents the results for $\lambda^* = 0.15$.[25] It should be mentioned that the critical values reported in Vogelsang (1997) are different when the errors are $I(0)$ in comparison to errors that are $I(1)$. In cases in which uncertainty persists with regard to errors being $I(0)$ or $I(1)$, the conservative approach of Vogelsang (1997) should be used, which recommends the use of critical values that corresponds to $I(1)$ errors. This is because when the errors are highly persistent, using this conservative approach provides a better finite sample approximation.

The results suggest that the dispersion of individual price changes experienced a

[25] The choice of λ^* is the minimum available trimming tabulated in Vogelsang (1997), given both the initial lag–length k chosen and sample size T of the Nicaraguan data.

structural break in its mean during the structural adjustment program initiated in April 1994, confirming our hypothesis based on the examination of this time series in Figure 3. All three statistics, *Sup*, *Mean*, and *Exp* reject the null of no structural break at the 5% significance level.[26]

Using the conservative approach explained above for the series that failed to reject the null of nonstationarity when (14) was estimated (which are $\{S\}$ and $\{I_X\}$), the hypothesis of a stable mean cannot be rejected. That is, using the critical values for $I(1)$ errors, which is the conservative approach when it is not known whether the errors are $I(0)$ or $I(1)$, there seems to be evidence of a stable mean for the time series $\{S_t\}$ and $\{I_{X,t}\}$ (in Table 4, the results reported are based on the conservative approach). When the critical values for $I(0)$ errors are used, these same series, skewness and inflation excluding food and energy, experience structural changes in the period. In the case of $\{S_t\}$, the *Mean* statistic rejects the null of no structural break at the 10% significance level, while the *Exp* and *Sup* statistics reject the null at 5% significance level. The estimated dates of the breaks are 1995:06 and 1993:12 respectively. Finally, note that in Table 4, inflation given by the α–trimmed means are not reported. But since inflation in the CPI does not display any instability in its mean, these measures, by construction, will not display any breaks either.

Recall that unit root testing indicated that core inflation, given by I_X, was not

[26] The *Exp* statistic is designed to have power in detecting large breaks, while the *Mean* statistic is designed to have power in detecting small breaks.

stationary around its mean. But the possibility of a break in its mean may be biasing the parameter γ in (14) towards zero (see Table 3). That is, a regression like (14) may be mispecified. The final step, then, is to test for the unit root hypothesis allowing for the possible change in the mean of the series documented in Table 4. This is done following the procedure designed by Perron and Vogelsang (1992), which basically consists of estimating (13) but including a dummy variable that takes the value of zero up to the estimated break date and one afterwards. The estimated coefficient for γ is also reported in Table 4 under the column t_γ. In the case of dispersion of relative prices $\{SD_t\}$, the hypothesis of nonstationarity is rejected at the 1% significance level, but for the case of skewness, $\{S_t\}$, the statistic fails to reject the null of nonstationarity at all significance levels. In the case of inflation excluding food and energy, $\{I_{X,t}\}$, the results are mixed since the null of nonstationarity is rejected at the 10% but not at the 5% significance level.[27]

The above results indicate that whereas the dispersion of relative prices, $\{SD_t\}$, experiences a structural decline in its mean in 1994:04, the time series for skewness, $\{S_t\}$, seem to have a stable trend function, while there are mixed results in the case of inflation excluding food and energy $\{I_{X,t}\}$. However, from Table 1, it is clear that, at least in the case of skewness, the errors may not be $I(1)$ but highly persistent. In this case the unit root test lacks the power to reject $I(1)$ errors, and the skewness

[27] The asymptotic critical values for this test were obtained from Tim Vogelsang directly, and appeared in a working paper version of Perron and Vogelsang (1992). These critical values are, at the 1%, 5%, and 10% significance levels, respectively: -4.945706, -4.432140, and -4.182082.

series may be stationary, but may have experienced a shift in its mean.

Before ending this section, it will be useful to test some of the relationships that have been widely tested in the literature, namely whether general inflation affects the dispersion and skewness of individual price changes. Following Lack and Tsiddon (1992), we test two models for the inflation-dispersion relationship. The first model or Model 1 is given by

$$SD_t = \phi_0 + \phi_1 \Pi_t + \phi_2 DU_t + u_t \tag{15}$$

where DU_t is a dummy variable that takes the value of 0 up to 1994:04 and 1 afterwards, and u_t is a disturbance that does not preclude the possibility of serial correlation. The inclusion of the dummy variable in (15) serves to account for the possible break in the mean dispersion of relative prices that occurred in 1994:04, when comprehensive structural reforms took place (see the results in Section 5, and in particular Table 4). For completeness Model 1 is also estimated without the dummy variable and the results are reported in Table 5.

Table 5 also presents results for an alternative model, Model 2, namely

$$SD_t = \phi_0 + \phi_1 E\Pi_t + \phi_2 U\Pi_t + \phi_3 DU_t + u_t \tag{16}$$

where $E\Pi_t$ and $U\Pi_t$ are expected and unexpected inflation, DU_t is the dummy variable defined above, and u_t is the error term. Expected inflation is estimated in a *naive* manner, with a distributed lag model using the general to specific approach suggested by Hendry (1975). That is,

$$E\Pi_t = \widehat{\vartheta}_0 + \sum_{i=1}^{k} \widehat{\vartheta}_i \Pi_{t-i}$$

All tests for no serial correlation in this simple forecast specification cannot be rejected at all significance levels.[28] Table 5 clearly shows that relative price change dispersion is positively and significantly related with general price change variability, and although both expected and unexpected generalized inflation affect relative price change dispersion positively, only unexpected inflation is statistically significant. As in the case of Model 1, Model 2 is also estimated without the dummy variable.

Note that the inclusion of the dummy results in a better specified model. Notice that the Durbin Watson statistics are closer to 2 in the both of the cases where the dummy was included. In addition, both the Akaike and Schwartz information criteria select the models that include the dummy. Finally, although in both cases the Q–statistics fail to reject the hypothesis of serial correlation, the p–values are higher in the cases in which the dummies are included (Table 5 reports the Q–statistic for one lag, but similar patterns are found for the remaining lags [the latter are not reported]). In brief, during the implementation of the structural adjustment that took place in 1994:04, the monthly dispersion of price–relatives experiences a structural decline in its mean that is estimated to be about 1%.

[28] Coincidentally, using the Box–Jenkins model selection methodology, this model ranks better than other ARIMA specifications (based on the Akaike and Schwartz information criteria, after adjusting for the number of observations).

In addition, the results reported in Table 5 stand in sharp contrast with findings that the changes in relative price change dispersion associated with the variability in the price level is the result of rigidities in prices due to menu costs (Lach and Tsiddon, 1992; references therein), in which case one would expect the coefficient for expected inflation to be larger than that of unexpected inflation (and possibly significant). Rather, the results in Table 5 suggest that it may be imperfect information of the Lucas–type that is accounting for this relationship (Lucas, 1973; Cukierman, 1984).

The study of the relationship between skewness, $\{S_t\}$, and general inflation in a simple framework as in the case of relative–price dispersion and inflation lead to mispecification. One of the reasons is that the $\{S_t\}$ may be nonstationary or at least highly persistent (see Table 4). A regression like (15) in this case is mispecified, with the resulting errors being serially correlated.[29]

[29] In fact, when regressing S_t on Π_t, the Ljung–Box Q–statistics with 1, 3, 6, and 12 lags (prob. in parenthesis) are 6.7296 (0.009), 10.555 (0.015), 25.082 (0.001), and 27.593 (0.004), respectively, indicating that the hypothesis of no serial correlation (white noise) is rejected at the 1% significance level for almost all lags. The Breusch–Godfrey serial correlation LM–test at 3, 6, and 12 lags (prob. in parenthesis) are 2.8356 (0.0429), 4.2578 (0.0008), and 2.5786 (0.0063), respectively, also indicating serial correlation in the estimated errors. However, ARCH test show that for various lags, we cannot reject the hypothesis of no ARCH.

6 Ranking of measures and forecasting

There is no clear way of ranking the various measures of core inflation studied. However, if inflation targeting were adopted as the main policy, the monetary authority will have to be able to forecast inflation relatively well. Under inflation targeting, short–run as well as long–run targets have to be set in advance to guide policy. The results obtained in Section 3 indicate that for the Nicaraguan price data, the CPI, because of high frequency noise, is not the MLE of core inflation. The normality assumption on the distribution of the idiosyncratic component of prices is in general violated by the data in the short–run. But the results also indicate that in the long–run such deviations tend to decrease at lower frequencies. Thus, the CPI will be useful in the formulation of long–run policy targets. Therefore, good measures of core inflation should remain close to the CPI in the long–run, while able to correct for the high frequency noise in the short–run.

Given that the alternative measures of inflation correct to a certain extent for the short-run noise, it remains to evaluate how close these indicators predict long–run inflation given by the CPI. Table 6 presents test statistics that facilitate evaluation and comparison of the forecasting performance of the various measures of inflation. The summary statistics are the mean error (ME), the mean absolute error (MAE), and the root mean square error ($RMSE$). The procedure is as follows. First, using CPI inflation, a 24–month moving average is constructed, denoted I_{LR}.[30] Denote now

[30] A 36–month moving average was also used, but the results were unaltered (in relative terms).

$e_{jt} = I_{j,t} - I_{LR,t}$, where j stands for the measure used in computing e_{jt}. The mean error of measure j, which is given by, $ME_j = \frac{1}{T}\sum_{t=1}^{T} e_{jt}$, represents the average magnitude by which measure I_j differs from I_{LR}. The mean absolute error, given by $MAE = \frac{1}{T}\sum_{t=1}^{T} |e_{jt}|$, is a measure of accuracy, and the root mean square error, given by $RMSE = \left(\frac{1}{T}\sum_{t=1}^{T} e_{jt}^2\right)^{1/2}$, is an alternative measure of accuracy. According to ME all measures of underlying inflation tend to underpredict long–run inflation. In this respect, the various measures display considerable coherence. Under this criterion, the measures can be ranked from better to worse as follows: I_V, I_X, I_{15}, I_{50}, and I_{25}. With respect to accuracy, the 15%–trimmed mean, I_{15}, does unambiguously better (see Table 6). In this case the ranking is I_{15}, I_{50}, I_{25}, I_X, and I_V. When these statistics were computed for the post–adjustment period (1994:05-1998:12) the results were unchanged in relative terms.

The above results, along with the results obtained from the time series analysis, make a compelling case against the I_X measure. If it is the case that I_X has errors that are not $I(1)$ but very persistent, then a structural break in its mean might have occurred during the period. If this is the case, then the unit root test, after incorporating the possible break, may be lacking power in rejecting $I(1)$ errors. In any event, the measure I_X is not likely to forecast long-run inflation as well as other measures that have a higher correlation with CPI inflation, and that are less volatility,

By construction, the 24–centered moving average has the advantage of allowing for a larger number of observations.

significantly stationary, and that change overtime more gradually.

7 Conclusions and comments

The evidence presented in the previous sections can be summarized as follows. First, the distributions of individual price changes in Nicaragua are generally leptkurtic and skewed. Such non-normality warrants the study of alternative measures of inflation.

Second, standard tests of nonstationarity show that the alternative measure of inflation that exclude food and energy may be lacking a long-run relationship with headline inflation. When tests are performed, which correct for structural breaks that may be biasing the Dickey-Fuller statistics towards zero, the results are somewhat mixed. While it is the case that the unit root hypothesis may be rejected with sufficient confidence, the test for structural breaks indicates that there is not sufficient evidence to reject the null of a stable mean. This result constitutes a warning against the use of this indicator in an inflation targeting regime.

Third, while inflation is a stationary process, skewness seem to have highly persistent errors. A possible explanation is that breaks in the mean of the standard deviation of the individual price change distribution may induce deviations of mean inflation from population estimates, which will nurture movements in higher moments, inducing the mean-skewness correlation observed in the data.[31] This opens an avenue for future research: using multicountry data one could test whether the persistence

[31] See Bryan and Cecchetti (1999) who show that the mean-skewness correlation seems to be the product of a small sample bias, which may lead to incorrect inference.

of the errors in the skewness time-series is commonplace in various countries. This will shed new light on the price formation mechanism.

Fourth, the dispersion of individual price changes seem to experience a structural break when the country adopted a structural adjustment program in April 1994, suggesting that such programs may contribute to a reduction in the dispersion of individual price changes. This has important implications regarding inflation, for it is also found that the reduction in relative price dispersion lead to a substantial, once and for all, reduction in mean monthly (headline) inflation.

Fifth, in terms of the forecasting ability of the various measures, it is found that all the indicators of core inflation tend to underestimate long-run inflation, but in terms of accuracy or efficiency, the 15% trimmed mean does unambiguously better than the other alternative indicators.

Finally, when testing the inflation and relative-price dispersion relationship, the evidence suggests that it is unexpected inflation, rather than expected inflation, that is related to relative-price dispersion. Further investigation of such relationship using more disaggregated data (perhaps at the establishment level) and using forward looking estimates of expected inflation, should shed more light on the mechanisms of price and expectation formation in developing economies, in general, and in the Nicaraguan economy in particular.

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Table 1. Summary statistics for the moments of the price change distribution for the 27 components of the CPI. Sample: 1991:06-1998:12

	WSD_t					
Horizon (months)	1	3	6	12	24	36
Mean	0.0367	0.0095	0.0057	0.0034	0.0025	0.0022
St. Dev.	0.0211	0.0037	0.0018	0.0010	0.0004	0.0003
	WS_t					
Horizon (months)	1	3	6	12	24	36
Mean	0.3838	0.4085	0.5116	0.4474	0.2608	0.0638
St. Dev.	1.7322	1.7554	1.6221	1.2830	1.2677	0.7770
	WK_t					
Horizon (months)	1	3	6	12	24	36
Mean	6.6311	6.7137	6.6191	5.6892	5.7424	4.4833
St. Dev.	6.4350	6.9522	7.1113	3.6655	3.3599	1.0151

Table 2. Correlation Matrix for different measures of core inflation and other summary statistics

Correlation matrix

Sample: 1991:06-1998:12

	inflation in CPI	I_X	I_V	I_{15}	I_{25}	I_{50}
inflation in CPI	1.000000					
I_X	0.598658	1.000000				
I_V	0.884852	0.612883	1.000000			
I_{15}	0.675645	0.727965	0.784366	1.000000		
I_{25}	0.660170	0.708015	0.768244	0.984285	1.000000	
I_{50}	0.613288	0.646518	0.711420	0.915017	0.950314	1.000000

Summary statistics

Sample: 1991:06-1998:12

	inflation in CPI	I_X	I_V	I_{15}	I_{25}	I_{50}
Mean	0.009954	0.008370	0.008094	0.007523	0.003642	0.004889
Median	0.008448	0.006216	0.007192	0.006721	0.003182	0.003551
Stand. Dev.	0.014035	0.013099	0.012566	0.006108	0.003637	0.006969

Table 3. ADF tests for nonstationarity

	Estimated coefficient	Other statistics	
Series	γ^a	F-Stat[b]	R^2
CPI (in logs)	-0.029	3.147	0.364
	(-0.358)	(0.0014)	
SD_t	0.947***	67.374	0.441
	(-8.208)	(0.000001)	
S_t	-0.998	6.153	0.531
	-1.980	(0.00001)	
Inflation rate - CPI	-1.903***	8.359	0.548
	(-4.375)	(0.000001)	
Inflation rate - I_X	-0.968	5.634	0.564
	(-1.910)	(0.00001)	
Inflation rate - I_V	-1.683***	6.73	0.494
	(-4.221)	(0.00001)	
Inflation rate - I_{15}	-1.493***	5.704	0.453
	(-4.557)	(0.00004)	
Inflation rate - I_{25}	-1.466***	5.225	0.448
	(-4.538)	(0.000005)	
Inflation rate - I_{50}	-1.375***	5.48	0.443
	(-4.633)	(0.000006)	

Notes: For each regression, the lag-length was chosen according to the procedure in Perron (1990), and is explained in detail in the text. *, **, and *** denote significance at the 10%, 5%, and 1% levels respectively.

a: The point estimate is $\gamma = \alpha$-1. In parenthesis, below the point estimate, is t_γ. Significance is based on MacKinnon critical values.

b: *p*-values in parenthesis.

Table 4. Wald type tests for detecting structural break

		Statistics				
Series	Sample	*Mean* [b]	*Exp* [b]	*Sup* [b]	*Estimated* T_b	t_g [c]
CPI	91/06-98/12	1.567	1.642	6.866	93/10	
SD_t	91/06-98/12	4.385**	2.684**	9.111**	94/04	-9.007***
S_t [a]	91/06-98/12	2.112	3.148	10.881	95/06	-3.096
Inflation in *CPI*	91/06-98/12	0.709	0.661	4.257	92/12	
Inflation in I_X [a]	91/06-98/12	2.355	3.134	11.242	93/12	-4.182*
Inflation in I_V	91/06-98/12	1.123	0.71	3.349	97/08	

Notes:

a: Since the critical values differ when the errors are $I(0)$ as compared to $I(1)$, the critical values used to evaluate the hypothesis of no break are those corresponding to $I(1)$. This follows the conservative approach suggested by Vogelsang (1997), see text.

b: *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively. Critical values are taken from Vogelsang (1997).

c: *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively. Critical Values are taken from Perron and Vogelsang (1992).

Table 5. Estimated effects of inflation, and expected and unexpected inflation, on the distribution of price changes

	Model 1		Model 2	
Regressors	Dummy	No Dummy	Dummy	No Dummy
Intercept	3.893***	2.870***	3.945***	3.040***
	(0.3698)	(0.269)	(0.5002)	(0.405)
P	0.701***	0.707***		
	(0.1468)	(0.157)		
$E\Pi$			0.325	0.335
			(0.3251)	(0.339)
$U\Pi$			0.679***	0.702***
			(0.1751)	(0.1826)
Dummy	-1.596***		-1.323***	
	(0.4252)		(0.467)	
R^2	0.3	0.19	0.25	0.17
Adj R^2	0.28	0.18	0.22	0.14
DW	1.88	1.72	1.992	1.93
AIC	1.362	1.49	1.318	1.47
SBC	1.445	1.56	1.437	1.54
F-Statistic	18.646	20.195	8.412	7.89
Prob.	0.000001	0.000021	0.000067	0.000076
Ljung-Box Q-stat (1 lags)	0.2633	1.4081	0.0003	0.1221
Prob	0.608	0.235	0.987	0.727
Ljung-Box Q-stat (3 lags)	1.2044		0.6182	
Prob	0.752		0.892	
Ljung-Box Q-stat (6 lags)	3.3003		2.496	
Prob	0.77		0.869	
Ljung-Box Q-stat (12 lags)	8.8742		6.5082	
Prob	0.714		0.888	

Note: For the regressors, standard errors in parenthesis. Significance at the 10%, 5%, and 1% level is denoted by *, **, ***, respectively.

Table 6: Forecasting performance of the core inflation measures

	Sample period: 1991:06-1998:12		
Measure of inflation	***ME***	***MAE***	***RMSE***
Inflation in CPI	-0.000150782	0.008970416	0.011822597
I_X	-0.00232266	0.00691482	0.00942791
I_V	-0.00209550	0.00805486	0.01005746
I_{15}	-0.00259876	0.003882516	0.00463115
I_{25}	-0.00611226	0.00624836	0.00663425
I_{50}	-0.00484807	0.00579725	0.00660727

Note: The forecasting performance is based on the ability of the different measures of core inflation to match a 24-month moving average of CPI monthly inflation.

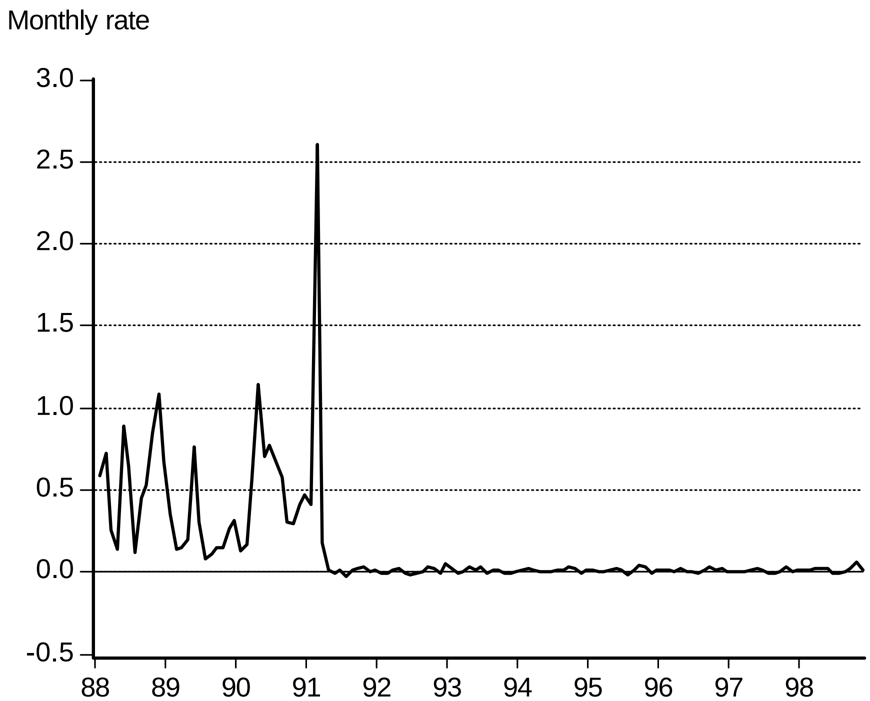

Figure 1. Monthly inflation in the CPI

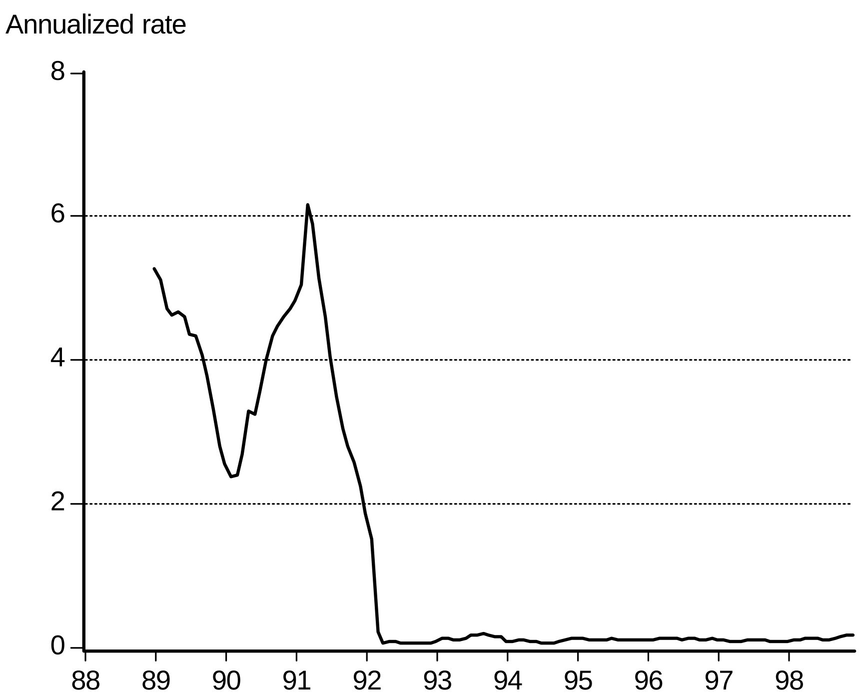

Figure 2. Annualized monthly inflation in the CPI

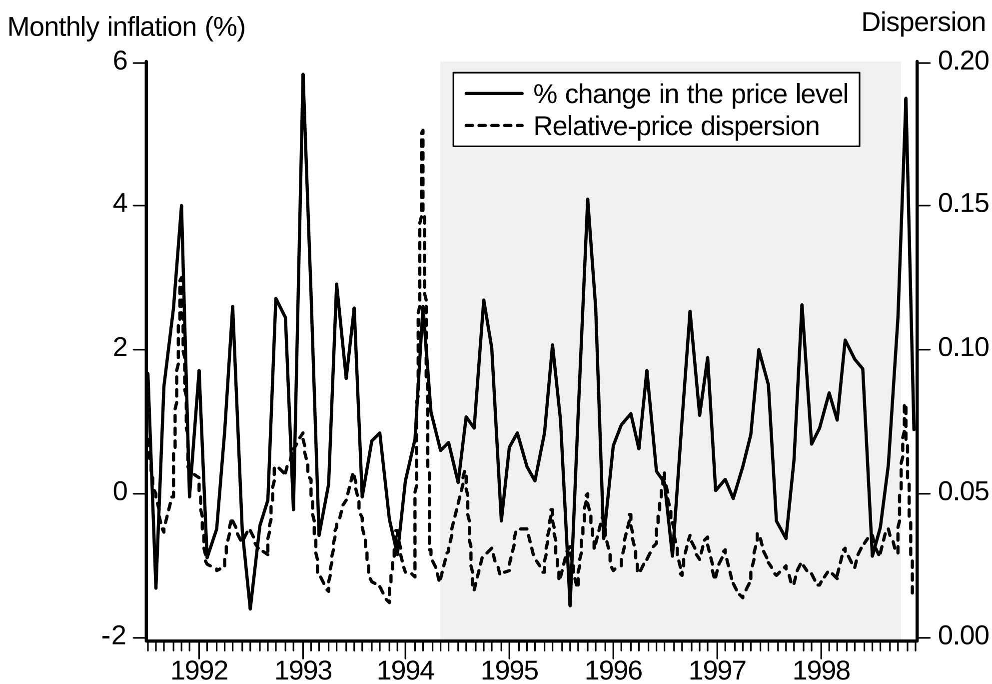

Figure 3: Monthly inflation in the CPI and relative-price dispersion

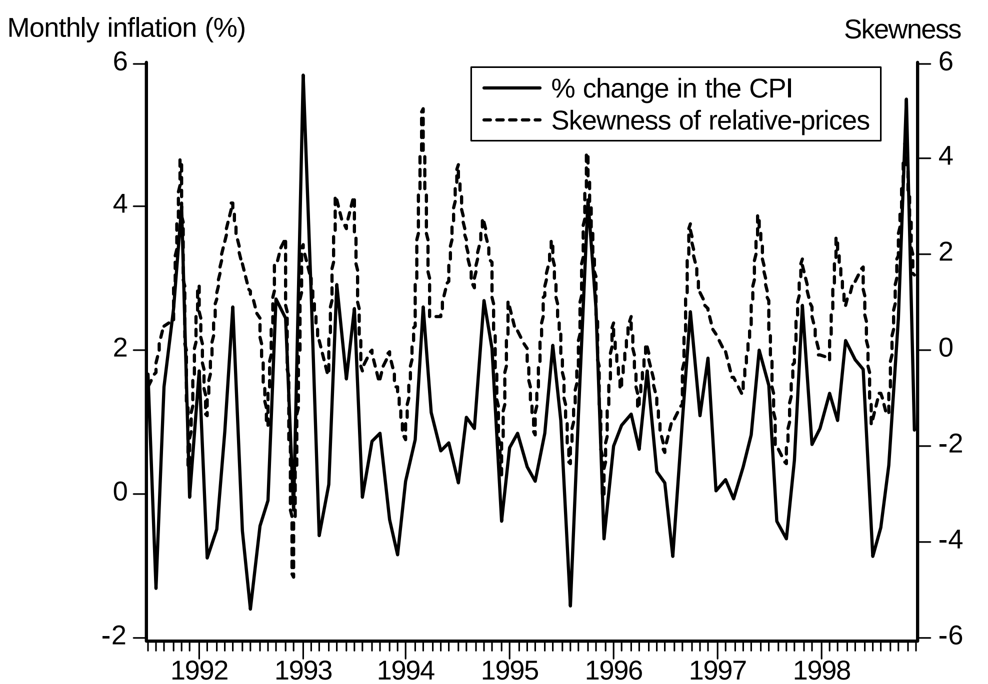

Figure 4: Monthly inflation in the CPI and skewness of relative-prices

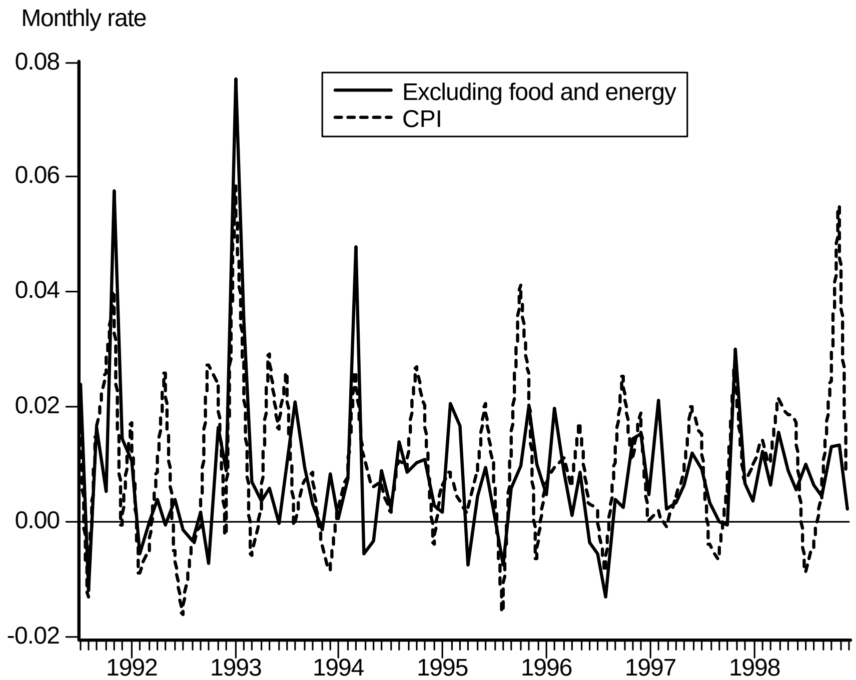

Figure 5: Core inflation excluding food and energy

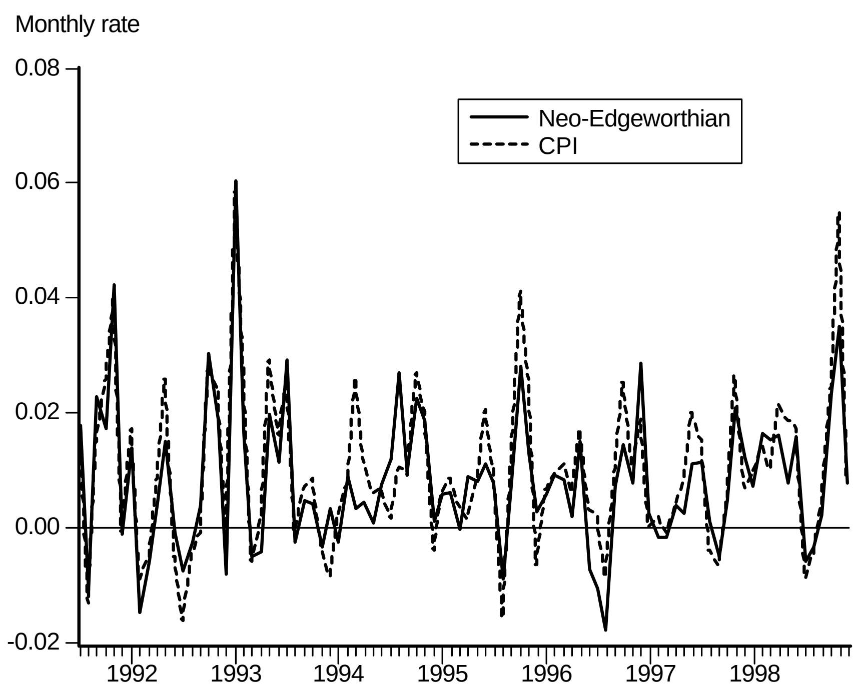

Figure 6: Core Inflation. The New-Edgeworthian measure

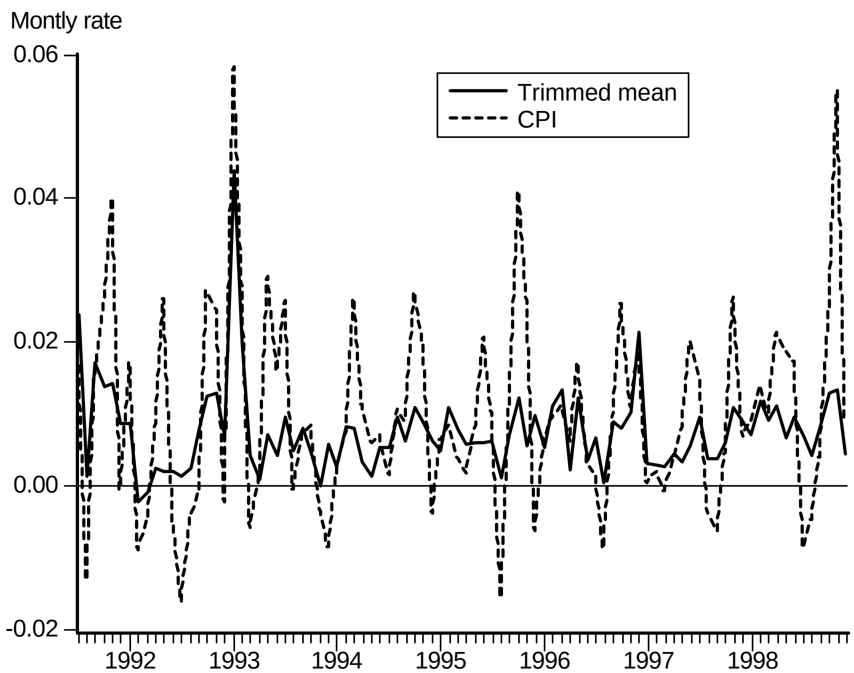

Figure 7: Core Inflation. 15% trimmed mean