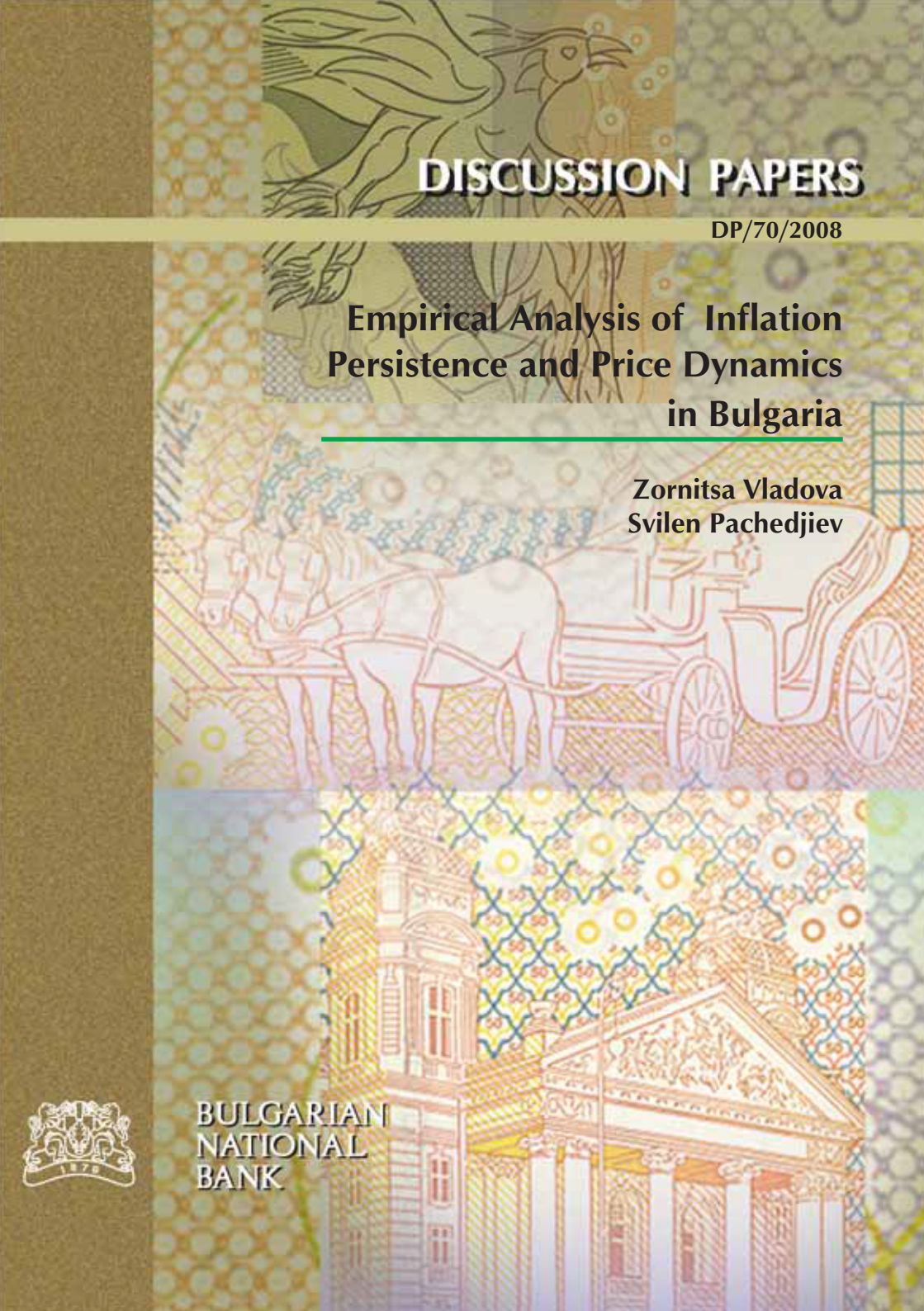
DISCUSSION PAPERS
DP/70/2008
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Zornitsa Vladova
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BULGARIAN NATIONAL BANK

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Send your comments and opinions to:
Publications Division
Bulgarian National Bank
1. Alexander Battenberg Square
1000 Sofia. Bulgaria
Tel.: (+359 2) 9145 1351, 9145 1978, 981 1391
Fax: (+359 2) 980 2425
e-mail: Dimova.L@bnbank.org
Website: www.bnb.bg

Contents

SUMMARY. The motivation of this paper stems from the extensive research on inflation persistence in euro area countries and the scant evidence on inflation inertia in Bulgaria. Persistence of inflation in Bulgaria can have implications for meeting the Maastricht criterion on inflation and its studying is also important for other reasons such as for forecasting purposes. We document the statistical properties of inflation in Bulgaria by producing persistence estimates for the headline HICP inflation, three measures of core inflation and the inflation of the main HICP subcomponents. The empirical analysis of persistence is based on the sum of autoregressive coefficients derived from unit root tests over the 1998–2007 period. The tests are applied to both constant-mean and de-trended inflation series measured as quarter-on-quarter rate of change and year-on-year rate of change at quarterly frequency. The results on persistence are discussed in the context of inflation persistence in the euro area and the new member states. We also analyse price dynamics in Bulgaria at representative products level over the 2000–2007 period and compare our conclusions with the evidence for euro area countries.

Key words: Inflation, inflation persistence, price dynamics

Zornitsa Vladova, Bulgarian National Bank, Economic Research and Projections Directorate, Vladova.Z@bnbank.org

Svilen Pachedjiev, Bulgarian National Bank, Economic Research and Projections Directorate, Pachedjiev.S@bnbank.org

1. Introduction

The main challenge Bulgaria faces on the road to euro adoption is the fulfillment of the Maastricht criterion on inflation. The achievement of the degree of nominal convergence related to price stability which is required for entering into the third phase of the Economic and Monetary Union raises various issues about the sources of inflation dynamics in Bulgaria. In recent years considerable attention has focused on the various factors driving inflation such as administered prices and excise tax rates adjustments, import prices, the Balassa – Samuelson effect, etc. However, empirical research on the nature or the behaviour of the inflation process itself, including a relative assessment of the similarities and differences between inflation in Bulgaria, the euro area and the other new member states (NMS) has not been undertaken so far.

One of the important characteristics of inflation which has been extensively analyzed for euro area countries is inflation persistence, that is the tendency of inflation to revert back slowly towards its long-run value following a shock which has moved it away from this long-run value. The analysis of inflation persistence for the euro area was conducted within a large-scale research project of the Eurosystem entitled *Inflation Persistence Network* which was implemented during the period of 2003–2005. The motivation for the project was derived from the crucial implications of inflation persistence for considering lags in the transmission of monetary policy, for possible country-specific effects of the common monetary policy and for the observed inflation differentials across euro area countries.

Assessing the degree of inflation persistence is of no less significance in the case of Bulgaria. Franta, Saxa and Šmídková (2007), among the few exploring persistence in some of the NMS along with that in the euro area, argue that a possible high level of national persistence could imply difficulties in meeting the Maastricht inflation requirement and eventual problems in sustaining convergence with the inflation rates of the other euro area members after euro adoption. Evaluation of the extent of inflation persistence in Bulgaria is important for a number of other reasons such as its relevance for forecasting purposes as well. To the best of our knowledge, however, the study of inflation inertia in Bulgaria on a comparable basis with the euro area (and the other NMS) has been limited so far. The present paper aims to fill this gap in the literature by documenting the statistical properties of the inflation process in Bulgaria and producing estimates of inflation persistence. This will constitute a useful initial step in the gathering of information on inflation inertia and in providing some preliminary evidence on this characteristic of inflation dynamics in Bulgaria.

The analysis of inflation persistence is carried out along the main directions of research conducted under the Inflation Persistence Network (IPN) for euro area countries with a view to enabling international comparison of the results. IPN has two general fields of research: macrolevel analysis and micro-level analysis.

The first field of research focuses on macroeconomic time series and evaluates the degree of inflation persistence at the euro area, country and sectoral level. This area of research is primarily concerned with obtaining purely statistical (time-series) or structural (multivariate) estimates of inflation persistence. The statistical form of estimates is based on univariate (autoregressive) models of inflation. These models are used to answer the question whether inflation is a random walk process (that is, a strongly dependent time series displaying permanent effects of shocks and thus returning only very slowly to its long-run value after a shock) or whether inflation is not persistent. Unlike univariate models where the sources of inflation inertia cannot be identified but only overall inflation dynamics properties are assessed, multivariate models enable the specifying of inflation persistence sources on the basis of structural parameters mainly within the framework of the New Keynesian Phillips curve.

Notwithstanding the undoubted power of structural models for policy analysis , univariate models have so far remained the predominant approach in the persistence literature due to their advantages of producing concise results across countries. In addition, when univariate models are carried out at a lower level of aggregation (across subindices), they can help investigate where the persistence properties of headline inflation come from. For this reason and as a starting point in inflation persistence analysis for Bulgaria, we decide to utilize the approach of statistical (time-series) estimates to measuring persistence and apply that approach at both the aggregate and the disaggregate level (i.e. for the main HICP subcomponents according to Eurostat classification). The main measure of inflation inertia we estimate is the "sum of autoregressive coefficients" which is by far the most widely used in inflation persistence studies. In estimating persistence, we also take into consideration one of the major findings of the IPN literature that results from univariate models are strongly dependent on the assumption of the mean of inflation which in the context of such models is seen as representing its equilibrium or long-run level. IPN results convincingly show that once changes in the mean of inflation are controlled for (either by introducing a break/breaks in the mean or by allowing for a time-varying mean), there is only moderate inertia in the quarter-on-quarter inflation in euro area countries. In our estimates of inflation persistence, we work with both

quarter-on-quarter inflation and annual inflation rates at quarterly frequency due to the importance of annual inflation for the Maastricht criterion. Our initial hypothesis is that a changing mean assumption might be a better representation of most of the inflation variables in Bulgaria than a constant mean- or a breaks-in-the-mean assumption. This is motivated by the fact that as a country undergoing a process of economic transformation, catching-up growth and real income and price convergence, Bulgaria is characterized by salient trends in various macroeconomic variables, including converging inflation rates.

The second field of research within the IPN project framework comprises analysis of microlevel data for consumer and producer prices. This analysis is based respectively on the individual price records used for the construction of the consumer price index and the individual information underlying the producer price index (PPI). One of the objectives of this second area of research is to provide quantitative evidence on price stickiness practices because possible price rigidity has implications for inflation dynamics and inflation inertia. This field of the IPN also includes firm-level surveys aimed at identifying key characteristics and determinants of the price setting behavior of firms. A study of micro consumer and producer data in Bulgaria which is also further extended by firm-level factors underlying price changes will certainly have the merit of disentangling some of the important factors driving inflation processes in our country. As such microlevel data and surveys are still not available, however, in this paper we analyze consumer price changes at the most disaggregate level possible (average monthly prices of representative products) and compare these results with those for the euro area. At this preliminary stage of analysis of price stickiness, producer prices are not included due to the lack of harmonization of the Bulgarian PPI with the respective indices of EU countries.

The rest of the paper is structured as follows. In Section 2 we provide definition of inflation persistence and methodological details on the approach used to produce the results. Section 3 gives an overview of the data variables and reports the results on inflation persistence. Section 4 discusses the estimates of inflation persistence in Bulgaria in the context of the IPN evidence. It also reports additional results on inflation persistence in Bulgaria, the euro area and the other NMS following a common methodological approach for all countries. Section 5 presents an analysis of price dynamics at representative products level based on average monthly prices for the country over the 2000–2007 period and discusses them in the context of IPN results. Section 6 summarizes the conclusions.

2. Definition of inflation persistence, standard methodological approach and implications for measuring persistence in Bulgaria

2.1. Definition of inflation persistence

The IPN project adopts the following definition of inflation persistence: "the tendency of inflation to converge slowly (or sluggishly) towards its long-run value following a shock which has led inflation away from its long-run value" (Altissimo, Ehrmann and Smets, 2006). In less formal terms, persistence of inflation refers to the length of time it takes inflation to return to equilibrium (baseline) after an unexpected change (Willis, 2003). As outlined in the extensive IPN literature, two major issues emerge with respect to this definition of inflation persistence: 1) the relatively low speed of inflation adjustment to a shock and 2) the assumption about the long-run equilibrium level or path inflation converges to following a shock or in other words, the assumption about the particular pattern for the baseline of inflation.[1]

In the context of the New Keynesian Phillips curve, the prevailing theoretical framework and specification for empirical analysis of inflation, there are three main sources of inflation inertia. The first one is *intrinsic* persistence which originates in past inflation and more specifically in the mechanisms of price- and wage-setting that are applying a backward-looking indexation rule. The second one is the so-called *extrinsic* persistence which stems from inertia in the drivers of inflation, e.g. from the persistent processes determining the output gap. The third one is the *expectations-based* persistence and it refers to the imperfect information available to private agents because of which their behavior necessarily involves learning and is not based on rational expectations capable of speeding up agents' responses to shocks hitting the economy.

2.2. Standard methodological approach

Two main approaches are employed for modeling inflation inertia in the literature. The first or classical approach, labeled as reduced-form Phillips curve, pursues to investigate the autocorrelation properties of inflation based on simple univariate autoregressive (AR) models. The primary focus of this approach is the question whether inflation possesses a unit root. In case it is

[1]The definition practically implies that the degree of inflation persistence shows the speed at which inflation responds to a shock. The faster the return to the equilibrium (long-run) level, the less persistent the inflation process is and *vice versa*.

found that the inflation time series follows a random walk, it is described as highly persistent because shocks have permanent effects on it and therefore the inflation process itself has no tendency to return to a long-run path. The second approach, structural multivariate estimates, deals with explicitly modeling the various components driving inflation persistence (intrinsic, extrinsic and expectations-based). With a view to measuring the general degree of inflation persistence in Bulgaria and making preliminary conclusions about the statistical properties of the inflation process itself, we adopt the classical approach and estimate univariate inflation models. Further and more extensive research in the vein of structural models for inflation in Bulgaria goes beyond the scope of the present paper.

The standard univariate AR process for representation of inflation series takes the following form (Andrews and Chen, 1994; Batini, 2002; O'Reilly and Whelan, 2004; Levin and Piger, 2004; Hondroyiannis and Lazaretou, 2004; Marques, 2004):

$$y_t = \alpha + \sum_{j=1}^{K} \beta_j y_{t-j} + \varepsilon_t \quad (1)$$

where y_t is the relevant inflation series at time t (usually measured as quarter-on-quarter inflation, seasonally adjusted or not seasonally adjusted) and ε_t is a random disturbance term (shock) with no serial autocorrelation but with possible heteroscedastic properties.

In terms of measuring inflation persistence or the autocorrelation properties of inflation series, equation (1) is usually presented with the following equivalent form:

$$\Delta y_t = \alpha + \sum_{j=1}^{K-1} \delta_j \Delta y_{t-j} + (\rho - 1) y_{t-1} + \varepsilon_t \quad (2)$$

where expression (2) has the following transformations of the variables given by (1):

$$\rho = \sum_{j=1}^{K} \beta_j \quad (3)$$

$$\delta_j = -\sum_{i=1+j}^{K} \beta_i \quad (4)$$

The parameter of interest with respect to the persistence of inflation is ρ which basically represents the sum of the autoregressive coefficients (SARC) in (1), thus indicating the serial correlation in the data. The crucial point about the SARC (the parameter of persistence) lies in the fact that it provides

important information regarding the degree of persistence of shocks to the inflation series and in this way it measures the sluggishness (or the speed) at which the inflation process responds to shocks. As Andrews and Chen (1994) put it in their influential paper, the long-run persistence properties of a time series are displayed by their impulse response function, with the latter being directly related to the SARC. Marques (2004) provides a detailed explanation of various other statistics used in the literature to measure inflation persistence. These statistics include the "spectrum at zero frequency", the "half-life" indicator and the "largest autoregressive root". Marques argues that the SARC is the best measure of inflation persistence. In fact, the SARC is also the most frequently employed measure for quantifying inflation inertia. For that reason, it is also the main measure of inflation persistence used in this paper.

As is well known, when $\rho=1$ the inflation series has a unit root (it is a random walk). The series is considered strongly dependent (highly persistent) because it is highly autocorrelated with its own lags and the contribution of temporary shocks is permanently built in it. Hence, the speed of the response to a given shock is very low and the series is incapable of going back to its equilibrium or long-run path.[2] It is thus straightforward to see that with a high persistence of inflation, in inflationary periods high inflation rates will be accompanied by high inflation in future periods, while in periods of low inflation highly persistent series will lead to falling inflation rate in the future. If $|\rho| < 1$, the inflation series is broadly characterized as stationary and as not strongly persistent because it returns back to its mean.

Even though expression (2) is nothing more than the familiar Augmented Dickey-Fuller (ADF) test which is used here to determine whether a time series is stationary or a unit root process, practical estimation of (2) and arriving at the precise degree of persistence for a given series is compounded by two types of difficulties. The first type stems from purely statistical (econometric) reasons, whereas the second is related to the definition of persistence and its actual implementation based on estimation of an equation such as that given by (2). Taking into account these difficulties, automatic application of the ADF test is generally considered as giving only naive estimates of persistence and is usually provided only as reference against more robust results. The various statistical and theoretical issues, which also

[2] In statistical terms, processes containing a unit root are represented by a flat sample autocorrelation function, with the impact of shocks to the series being permanent. In contrast, correlations in I(0) processes decline to zero at an exponential rate, implying the fact that all shocks have only a short-term effect on the given process.

have important implications for assessing persistence in Bulgaria, are well documented in the econometrics literature and the IPN studies for euro area countries. They will be briefly discussed below along with the possible solutions identified to overcome them and the relevant implications for Bulgaria.

The statistical problems associated with calculation of expression (2) are related to two main issues. The first one arises in the case when the autoregressive process has a unit root (i.e. when the persistence parameter $\rho = 1$) or when ρ is close to one. As Andrews (1993) explains, when ρ is large (close to or equal to unity), the estimate of ρ obtained from an ordinary least-squares (OLS) estimation of (2) has a downward bias (that is, it is generally lower than the true value and this is more so when estimation is done for a small sample of time series). The most common solutions established in the literature to correct for this downward bias include using exactly median-unbiased estimates of ρ as suggested by Andrews and Chen (1994) or the so-called Hansen's grid bootstrap procedure which also estimates median-unbiased ρ (Hansen, 1999). In practical terms, however, when simple OLS-estimates do happen to indicate relatively high persistence (usually when ρ>0.7 or ρ>0.75, see e.g. Gadzinski and Orlandi, 2004 and Lünnemann and Mathä, 2004), the above-mentioned procedures for overcoming the problem with the downward bias do not actually solve it; rather they only provide exact confirmation of the statistical information produced by the OLS estimation. In addition, Hansen's grid bootstrap procedure is rather computationally intensive. We finally make the important note that indeed, while the null (unit root) hypothesis of an ADF test as displayed by expression (2) is a point hypothesis, the alternative is an interval of parameter values. Therefore, any reference about the value of ρ derived indirectly from estimation of (2) with the ADF test only provides a general idea of the stationarity/nonstationarity of the respective series and as such should not be interpreted as the exact measure of inflation's persistence.

The second statistical issue refers to the choice of the autoregressive lag order K in (1) and hence of the lag order K-1 in expression (2). As we know, the lagged values of the dependent variable in (2) remove the serial correlation in the residuals. The choice of lag length is usually determined according to the Akaike (AIC) or Schwarz (BIC) information criteria. Many empirical IPN papers restrict the maximum lag order to 4 or 5 because of the usual quarterly periodicity of the data. Considering the fact that the number of lags used influences the point estimate of ρ and that different information criteria determine differing lag lengths, we conclude that the point estimates of ρ will become very sensitive to the number of lags utilized. Accordingly, it

may be more useful to consider both an ADF test based on automatic selection of lag length (e.g. according to AIC or BIC) and a Phillips-Perron (PP) test. Thus, estimation of (2) may be done with and without lagged differenced terms of the dependent variable.

When expression (2) is estimated with lagged differenced terms, it becomes a simple ADF test equation applied to a given inflation time series (y_t) where *K-1* can be automatically selected, e.g. according to BIC (max length=5) and the variable of interest to estimate indirectly is ρ.

Estimating (2) without lagged values of differenced inflation, we obtain

$$\Delta y_t = \alpha + (\rho - 1) y_{t-1} + \varepsilon_t \qquad (5)$$

which represents a PP test applied to the given inflation time series (y_t), for example employing the default options of EViews for spectral estimation method and bandwidth. Under expression (5) the main variable of interest is again ρ. We note that expression (5) does not follow directly from (2) and that in fact (5) is a modified version of (1) when the lag order *K* is set equal to 1. In that case (1) represents an AR(1) process and consequently, ρ is a measure of the lag one autocorrelation of the given inflation process or the AR(1) coefficient. The AR(1) coefficient is another measure of inflation persistence used in the literature (Altissimo, Mojon and Zaffaroni, 2004; Benczúr and Rátfai, 2005).

The second group of problems when gauging inflation persistence from equations such as (2) and (5) is that these equations/tests rely on the assumption of time-invariant coefficients. The assumption of constant coefficients may not be robust to occurrence of structural changes in the mean of inflation (i.e. its long-run value implied by the constant α) or to possible changes in the persistence itself (i.e. when there are changes in the autocorrelation properties of the inflation series over time concerning either the persistence parameter ρ or the other autoregressive parameters). From a research point of view based on the definition of persistence (as given in the beginning of the section) and the objective of obtaining estimates of the degree of persistence especially over long time periods as is the case of most IPN studies, this group of problems is probably more important than the statistical issues. The most often applied procedure to ensuring the robustness of results is allowing for changes in the mean of inflation and testing the stability of the persistence parameter ρ and the overall stability of the other autoregressive coefficients.

The predominant approach for considering possible changes in the long-run mean of inflation is to account for structural break(s) especially in case of

a change in the monetary policy regime or of other significant events which bring about permanent changes in the mean of inflation, e.g. VAT changes, administered prices adjustments, changes in the measurement of the consumer price index, etc. Statistically, in justifying the need to control for a break or breaks in the mean so as not to obtain overestimated estimates of inflation persistence, most authors draw on Perron (1989). Usually, the econometric techniques to identify structural breaks in the mean of inflation include classical hypothesis tests for breaks in the intercept both when the breakpoint date is known or presumed and when search over all potential breakpoints is made. Less often, Bayesian methods are employed. One of the main conclusions in the IPN literature regarding the treatment of structural changes in the long-run level of inflation is that estimates of inflation persistence become relatively lower when such changes are controlled for (e.g. Levin and Piger, 2004, Gadzinski and Orlandi, 2004, Lünnemann and Mathä, 2004, Corvoisier and Mojon, 2005, Cecchettti and Debelle, 2005). Testing for the presence of structural breaks in the mean of inflation series and allowing for such breaks in equations (2) and (5) is obviously not possible if these equations are estimated with the ADF and PP tests as the latter are automatically run in statistical software packages. To be able to control for a change or changes in the intercept in these equations, most researchers usually do not apply unit root tests directly. The most common approach adopted is to estimate (2) directly with OLS but not based on the ADF test. This allows extension of the equations with permanent shift dummy variables once formal tests prove the existence of such breaks in the given inflation series.

In perhaps one of the most influential IPN papers, Marques (2004) departs from the common approach of controlling for changes in the mean of inflation once structural break or breaks are identified. Arguing that "any estimate of persistence is conditional on the given assumption for the long-run equilibrium level of inflation", he stresses that the key point in assessing the persistence of inflation is actually the persistence of the deviation of inflation from its mean. In his opinion, models assuming an exogenous time-varying long-run mean of inflation are much more plausible than those including one or more breaks in the intercept especially when constant inflation mean or discrete shifts in the mean cannot be assumed for the economic development of a country. For this reason, he suggests that the general autoregressive model in (1) should refer to the demeaned (detrended) inflation series. Consequently, equations (2) and (5) should have as the dependent variable not differenced inflation but the differenced deviation of inflation from a possibly time-varying mean (proxied by filter-based trend components, moving-average methods or simple time trends). In

line with the studies obtaining low estimates of persistence when allowing for structural breaks in the mean of inflation, Marques also shows that persistence could be substantially lower under the assumption of a time-varying inflation mean.

Concerning the issue of changes in inflation persistence over time, the standard methods in the literature include: 1) splitting the sample into shorter periods of time and testing hypothesis about changes in the persistence parameter ρ; and 2) implementing rolling regressions in which calculations are done over sequences of short rolling samples.[3]

In sum, the standard methodological approach for assessing inflation persistence is based on univariate (AR) models of inflation with OLS estimation of the sum of autoregressive coefficients. Several critical issues need to be addressed in such models:

- the likely presence of downward bias in the persistence parameter when the latter is unity or close to unity;

- the appropriate choice of lag length;

- the need to avoid overestimation of persistence by allowing for changes in the mean of inflation (either by controlling for structural breaks or by considering a time-varying mean);

- the possible change of persistence over time especially over long time periods.

2.3. Characteristics of measuring inflation persistence in Bulgaria

In the practical measurement of inflation persistence for Bulgaria, we produce preliminary results based on OLS estimation by the ADF and PP tests of equations (2) and (5). For robustness of our results, we follow Marques's approach of assuming a time-varying mean of inflation rather than testing for possible structural breaks in the inflation series on the basis of the following arguments.

First, we have a relatively short sample (1998:q1 – 2007:q4) with no changes in the policy regime after the implementation of the currency board

[3]The method of rolling regressions is applied by O'Reilly and Whelan (2004), Gadzinski and Orlandi (2004), Hondroyiannis and Lazaretou (2004) and Pivetta and Reis (2006). As Hondroyiannis and Lazaretou (2004) point out, rolling regressions are another way of assessing the impact of possible structural breaks in the persistence parameter and in the autoregressive coefficients in an expression such as (2). In rolling regressions, the coefficients are estimated for a certain period of fixed-length which is then sequentially shortened and further extended forward, thus enabling to draw conclusions about the stability of parameter estimates.

arrangement. Even though there have been significant price adjustments over this period (especially concerning administered prices and excise tax rates) accompanied by possible shifts in the level of inflation series, controlling for all possible breaks is virtually hindered by the finite sample we have. To address the issue of structural break or breaks in the series, we decide to work with inflation series excluding administered prices (both at the aggregate and at the disaggregate level).

Second, the transformation period after the adoption of the currency board has been characterized by stabilization of inflation rates following the high inflationary period of 1997 and subsequent process of catching-up growth and price level convergence. In essence, these processes imply that an assumption of a time-varying mean, possibly medium- rather than long-term, is more appropriate for Bulgarian inflation series.

Third, according to a study exploring inflation persistence in few of the NMS (Franta, Saxa and Šmídková, 2007) time-varying mean models should be a preferred option for inflation persistence measurement in these countries. Even though that study employs structural rather than univariate autoregressive models, we make use of its final conclusion that inflation persistence in transition (NMS) countries should necessarily be assessed under a time-varying mean assumption.

On account of the relatively short inflation time series, we consider the issue of possible variation in persistence (e.g time-variant values of ρ) less relevant and do not investigate it empirically at this stage of analysis. In fact, our short time period precludes us from employing some of the standard methods for evaluating inflation's statistical persistence over time as discussed in Section 1.2.

Consequently, as we do not need additional flexibility in assessing the time variance of the coefficients in (2) and (5) or in allowing for structural breaks in the intercept in them, we can apply the standard ADF and PP tests for these equations. The persistence parameter of inflation can be derived indirectly from the results of the two unit root tests.

It should be noted that the use of ADF and PP tests based respectively on expressions (2) and (5) with a time-varying mean actually implies testing the persistence of the demeaned (de-trended) inflation series. In practical terms, this amounts to applying the two unit root tests without an intercept as shown below:

$$\Delta z_t = \sum_{j=1}^{K-1} \delta_j \Delta z_{t-j} + (\rho - 1) z_{t-1} + \varepsilon_t \qquad (6)$$

where (6) represents a modification of (2) with the following two changes: y_t (the inflation series of interest) is replaced by z_t (the series of the deviations of inflation y_t from its time-varying mean μ_t, i.e. $z_t = y_t - \mu_t$); the intercept α in (2) has been removed.

$$\Delta z_t = (\rho - 1) z_{t-1} + \varepsilon_t \qquad (7)$$

where as with (6), equation (7) is a modification of (5) with the following changes: y_t is replaced by z_t (the series of the deviations of inflation y_t from its time-varying mean μ_t); the intercept α in (5) has been removed.

3. Presentation of inflation data and estimates of univariate autoregressive models

3.1. Data description

The most common type of inflation rate employed in the literature to assess inflation persistence is the quarter-on-quarter inflation. We also use quarterly data of the consumer price index for Bulgaria (CPI/HICP). Our timespan covers the period 1998:q1 – 2007:q4. At the aggregate level, we work with the overall index of consumer prices (CPI/HICP) and as proxies for core inflation we use three alternative indicators: HICP excluding administered prices and tobacco (HICPXADT), HICP excluding energy and unprocessed food (HICPXEUN) and HICP excluding energy, food, administered prices and tobacco (HICPXEFADT). Additionally, we investigate the persistence properties of the overall inflation of administered prices and tobacco (ADT). At the disaggregate (sectoral) level we work with HICP energy excluding administered prices (HICPEN), HICP processed food excluding tobacco (HICPFPR), HICP unprocessed food (HICPFU), HICP non-energy industrial goods excluding administered prices (HICPIG) and HICP services excluding administered prices (HICPSERV).[4] To get an idea about the persistence of the underlying or core inflation in the services sector, we construct an additional subindex of services which excludes administered prices, catering and transport services (HICPSERVXADCT).[5]

[4]This decomposition of the consumer price index corresponds to the Eurostat classification which presents in a comparable manner the major components of the HICP for the euro area and the EU countries. For the purposes of this paper and with a view to disentangling the effects of adjustments in administered prices and excise tax rates, we exclude administered prices and tobacco products from the main HICP components according to the Eurostat classification.

[5]Constructing such a measure of core services inflation, we aim to exclude the impact of administratively set prices (e.g. heat energy, refuse collection, sewerage collection, postal services, hospital services, etc.) and the impact of those components of services which are directly linked to the volatile food and crude oil prices.

Quarterly inflation rates are calculated following the standard way as the annualized quarter-on-quarter rate of change of the respective consumer price index.[6] As there is no established treatment of the issue of seasonality when inflation persistence is evaluated, we work with both non-seasonalized inflation rates and seasonally-adjusted rates.[7]

Due to the importance of assessing inflation persistence of year-on-year inflation rates, we also calculate annual inflation rates based on the respective quarterly indices.[8]

Figure 1 in the Appendix presents the annualized quarterly inflation rates (not seasonally-adjusted). We observe that the predominant part of inflation series have a mean-reverting behaviour, hence no substantial persistence is present in them.

Our initial expectations about trending (time-varying mean) inflation rates are confirmed for the series excluding the most volatile components of the HICP. These are processed food, non-energy industrial goods, services excluding administered prices, catering and transport and respectively core HICP inflation (excluding energy, food, administered prices and tobacco). We note that for these series an assumption about a constant mean over the whole period is not plausible.

More specifically, processed food quarterly inflation shows an upward trend over the last three years (from 2005 to 2007). Similarly, a clear acceleration of inflation has been evident for non-energy industrial goods excluding administered prices since 2004. For services quarterly inflation the period until the middle of 2002 is a period of disinflation and this might be reflecting the macrostabilization of the economy following the introduction of the currency board. The period from 2003 to 2004 is not characterized by any changes in the mean of inflation for the two groups of services. From the beginning of 2005, however, different patterns are observed. While services (HICPSERV) inflation rate starts accelerating on account of the catering and transport services, services inflation excluding these components speeds up less significantly. Accordingly, the core HICP inflation (excluding energy, food, administered prices and tobacco) follows the general trends of non-energy industrial goods and services inflation, showing acceleration since the beginning of 2005.

[6]The following standard transformation is applied: $\pi_t = 400*ln(Pt/ Pt\text{-}1)$, where Pt is the respective consumer price index at the aggregate and disaggregate (sectoral) level, 2005=100.

[7]Inflation rates are deseasonalized with Census X12-procedure implemented in EViews, using additive adjustment.

[8]Persistence in yearly inflation is not analyzed for the euro area and the other NMS countries and thus our results are not comparable.

It is worthwhile noting that the all conclusions made above based on visual inspection of the quarterly inflation series refer to inflation which contains seasonal variation. Deseasonalized series may be expected to exhibit a higher and/or different degree of persistence than the non-seasonally adjusted ones.

Practical estimation of inflation persistence is carried out in two steps. Initially, we estimate (2) and (5) at the aggregate and disaggregate level under the hypothesis that mean inflation is constant throughout the whole period (both for the non-seasonally adjusted and the seasonally-adjusted series). These estimations serve as reference for the second stage of analysis which considers mean inflation as a time-varying process (equations 6 and 7). Comparison of the results from the two stages allows evaluation of the impact of a changing mean of inflation on persistence estimates.

To define the time-variant mean level of inflation, we apply Hodrick-Prescott (HP) filter and do not resort to other statistical techniques for extracting the underlying trend of a series such as moving average methods and time trends. It is assumed that the use of moving-average methods (e.g. 13 terms centered moving average) provides a rather flexible mean level (long-run trend) of inflation and this may erroneously result in an underestimated degree of persistence (Marques, 2004). On the other hand, the use of time trends is not applicable since such trends are not present in a clear-cut way in the data of quarterly inflation series over the whole period.

Figure 2 in the Appendix, displaying year-on-year inflation rates, shows that a time-variant mean inflation should also be the primary assumption in gauging the persistence of these inflation series. Hence, our main results for the year-on-year inflation rates will be derived from (6) and (7) too. Again, as for the quarter-on-quarter inflation rates, we employ the HP filter-based trend component as a proxy for the time-varying mean of inflation.

3.2. Persistence estimates for quarter-on-quarter inflation

In this part we report the results on the persistence of quarterly inflation based on the ADF and PP unit root tests as specified in equations (2), (5), (6) and (7). First, we perform estimation of the persistence of quarterly inflation for the aggregate consumer price index (CPI/HICP) and then empirically examine the persistence properties of the more disaggregate groupings of the index.

Tables 1 and 2 in the Appendix show the values of ρ inferred indirectly from the two unit root tests, respectively for the non-seasonally adjusted and

the seasonally adjusted quarterly inflation rates. Table 1 displays the estimates of persistence obtained for the non-seasonally adjusted inflation rates based on the assumption of a constant mean (i.e. equations 2 and 5). As can be seen, at the aggregate level the two tests reject the null hypothesis of non-stationarity in all cases at the 1 per cent level. In addition, since the persistence parameter ρ does not exceed 0.35 we can conclude that a potential downward bias on the obtained estimates is not a particular problem. These results point to an overall low level of inflation inertia. It is noteworthy that all three measures of core inflation appear to be relatively more persistent than headline inflation, with HICPXEFADT, comprising services and non-energy industrial goods excluding administered prices, having the highest degree of persistence (ρ=0.35). This can be explained by the fact that HICPXEFADT excludes the most volatile components such as energy and unprocessed food and consequently its inflation rate becomes comparatively more inertial. Administered prices and tobacco (ADT) inflation also seems to lack persistence which might be due to the relative volatility of this series (ρ=0.07). The exclusion of ADT from HICPXEFADT is another factor for the higher relative persistence of HICPXEFADT.

At the sectoral level, HICP energy and HICP unprocessed food are the least persistent inflation series as they display no positive serial correlation. Turning to the other HICP subcomponents, we find that the unit root hypothesis is rejected for all but one subgroups (according to the ADF test), while the PP test rejects unit root processes for all cases. According to the PP test, processed food, non-energy industrial goods and both services inflation series have a relatively small persistence parameter varying between 0.4 and 0.55. The ADF test does not reject the unit root hypothesis at the conventional significance levels for non-energy industrial goods inflation series and for this variable the BIC criterion selects 5 lags, indicating the presence of significant correlation in the series.

Running the ADF and PP tests for the detrended quarterly inflation rates, we obtain considerably lower persistence estimates at both the aggregate and disaggregate level than from the constant mean estimates (Table 1A). This finding is expected considering the already observed varying trending pattern in some of the quarterly inflation variables (processed food, non-energy industrial goods, services, services excluding catering and transport ones and the core HICP excluding energy, food, administered prices and tobacco). An intuitive way to evaluate the impact of the changing mean of inflation on persistence estimates is to compare persistence parameters for the two cases: constant mean and time-varying mean (cf. Table 1 and Table 1A). Unsurprisingly, the biggest downward impact is observed for the inflation rates of industrial goods and services and for the core inflation

variables associated with them (HICP excluding energy and unprocessed food – HICPXEUN and HICP excluding energy, food, administered prices and tobacco – HICPXEFADT).

As expected, with seasonally-adjusted inflation rates, persistence estimates are somewhat higher than those for non-deseasonalized data because the volatility (variability) arising from seasonal variation has been removed (Table 2 and Table 2A). Nevertheless, the constant mean assumption estimates show that the null non-stationarity hypothesis can be rejected for all but one inflation series. Again, non-energy industrial goods inflation exhibits the highest level of persistence, with ρ reaching 0.7 and in that case, the two tests fail to reject the null hypothesis. In the case of processed food, both tests appear to provide weak statistical evidence against the random walk hypothesis as the latter is rejected at the 10 per cent level. Our hypothesis is that the constant mean assumption underlying the results for these two series might not be a good representation of the trending behaviour observed in them.

As Table 2A shows, all deseasonalized and de-trended inflation series are characterized by a low degree of inflation persistence which does not exceed 0.44. Thus, not only do we have strong evidence against high persistence in all inflation series, but we are also aware that a possible under-estimation of the persistence is not an issue. Persistence estimates produced by the ADF and the PP test are identical as for all indices a lag length of one quarter (i.e. $K=1$) is the predominant outcome. In Table 3 we compare the estimates of ρ obtained following the time-invariant and time-varying mean approaches.

Table 3. Comparison of inflation persistence estimates
Constant mean (CM) vs. time-varying mean (TVM) assumption
Annualized quarter-on-quarter inflation rates, seasonally adjusted
Sample period: 1998:q1 - 2007:q4

Inflation variable	ρ		difference
aggregate level	CM	TVM	ρ(CM)-ρ(TVM)
CPI/HICP	**0.24**	**0.13**	0.11
HICPXADT	**0.36**	**0.18**	0.18
HICPXEUN	**0.52**	**0.40**	0.12
HICPXEFADT	**0.47**	**0.21**	0.26
ADT	**0.07**	**-0.14**	0.21
disaggregate level			
HICPEN	**-0.03**	**-0.06**	0.03
HICPFPR	**0.62**	**0.40**	0.22
HICPFU	**0.01**	**-0.17**	0.18
HICPIG	0.74	**0.44**	0.30
HICPSERV	**0.66**	**0.15**	0.51
HICPSERVXADCT	**0.67**	**0.14**	0.53

Notes:
Inflation series are denoted as given on page 14.
They are seasonally-adjusted with Census-X12 procedure, using additive adjustment.
ADT (administered prices and tobacco) inflation is not seasonally adjusted.
The estimates of inflation persistence for ADT are the same as those obtained from the non seasonally-adjusted data.
The estimates of ρ are obtained indirectly from the coefficient on (p-1) produced with OLS from the ADF test.
Numbers in bold indicate that it is possible to reject statistically that ρ=1
See Table 1 and Table 2 in the Appendix for more details.

Services, non-energy industrial goods and the core HICPXEFADT inflation variable have the biggest downward impact on their inflation persistence once the latter is estimated for de-trended inflation series. This is an indirect evidence of the presence of a pronounced trending pattern in these inflation variables which should be taken into consideration when measuring their persistence. Figures 3 through 6 illustrate this point. HICP services and HICP services excluding catering and transport ones (HICPSERVXADCT) inflation variables share a downward trend until 2002 and a rather stable rate of change from 2003 to 2004; this corresponds to our preliminary observations in Section 2.1. Following 2004, however, HICP services inflation has an upward trend, whereas the growth pattern of HICPSERVXADCT is much less prominent. This shows that developments in overall services inflation (excluding administered prices) are primarily driven by catering and transport services. Some factors which may account for this are: 1) the larger share of

the two groups in the consumer basket since 2005 (with the introduction of the HICP methodology of compilation of the consumer price index); 2) the stronger impact of cost-push and structural factors underlying catering and transport services inflation (e.g. international oil prices hikes, weather conditions, more pronounced price level convergence concerning food prices and catering, etc.).

Figure 3

SERVICES EXCL. ADMINISTERED PRICES INFLATION

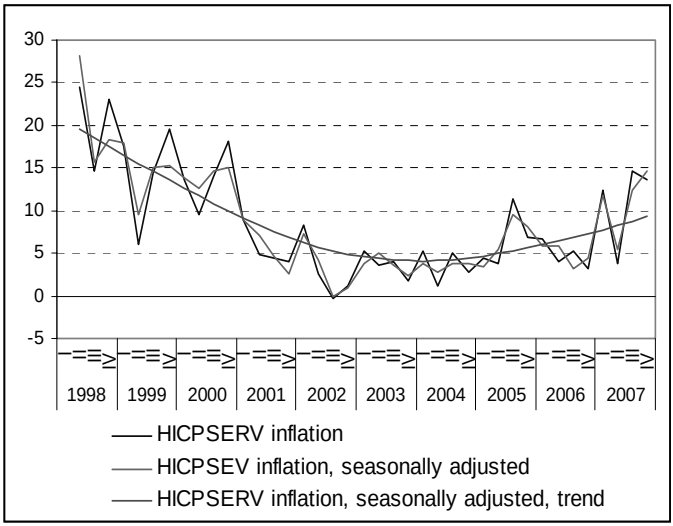


Figure 4

SERVICES EXCL. ADMINISTERED PRICES, CATERING AND TRANSPORT INFLATION

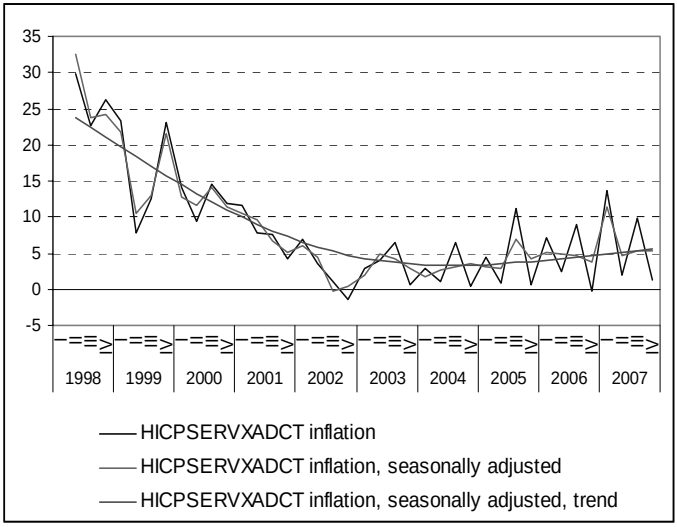


Figure 5

NON-ENERGY INDUSTRIAL GOODS EXCL. ADMINISTERED PRICES INFLATION

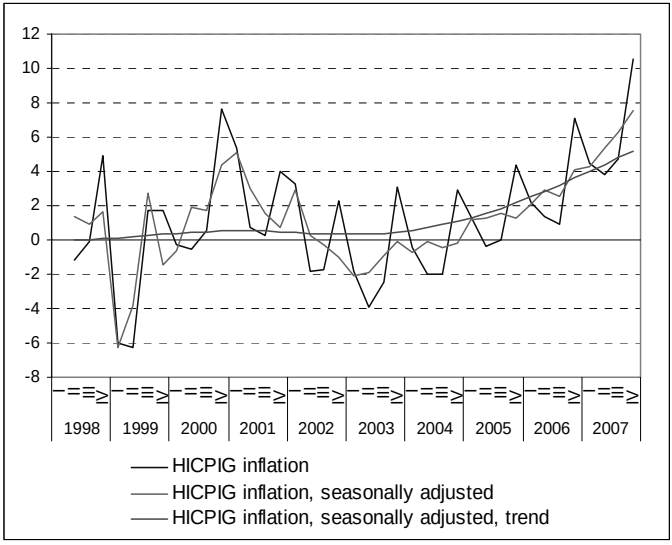
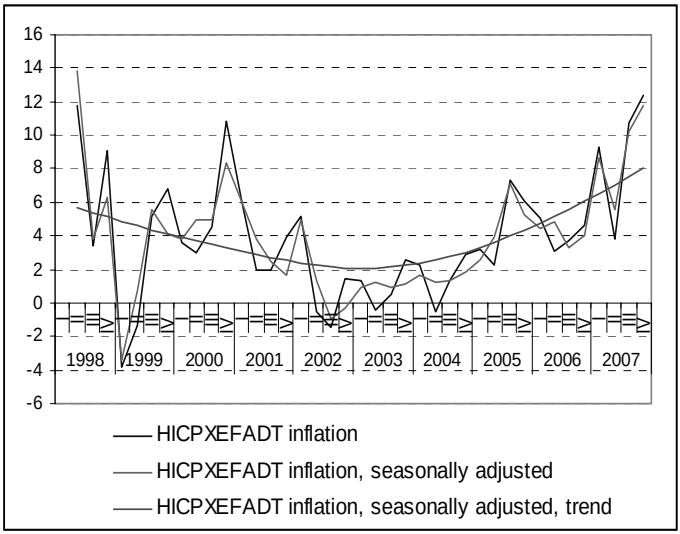


Figure 6

HICP EXCL. ENERGY, FOOD, ADMINISTERED PRICES AND TOBACCO INFLATION



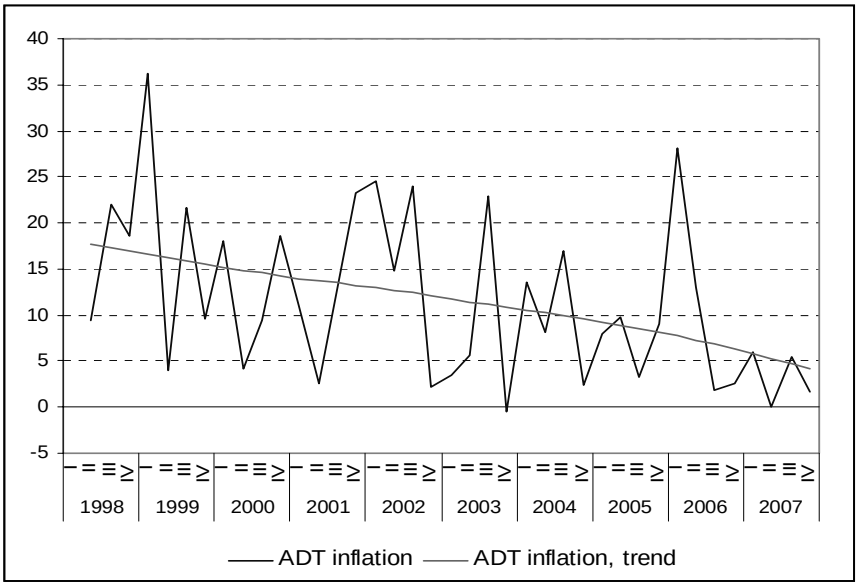
Similarly, non-energy industrial goods (HICPIG) inflation has also been accelerating since 2005, even though from much lower initial levels. Inflation developments in this group over the last 3 years may be explained by the increased consumer spending associated with rising real disposable incomes and the facilitated and cheaper access to various forms of consumer credits made available with the enhanced competition in the financial sector. Another factor that might explain the acceleration in non-energy industrial goods inflation is the gradual improvement in the quality of these products with the ongoing process of meeting EU standards. Driven by the observed inflation processes in services and non-energy industrial goods, core HICP excluding energy, food, administered prices and tobacco (HICPXEFADT) inflation has largely followed them with a clear-cut upward trend since 2005.

Going back to Table 3 where we explicitly allow for the trending processes in the seasonally-adjusted inflation rates, we can summarize the following conclusions on inflation persistence in Bulgaria. At the aggregate level, headline CPI/HICP inflation is characterized by an overall low degree of persistence (ρ=0.13) mainly due to the negative serial correlation evident for unprocessed food inflation (ρ=-0.17) and to a lesser extent that for energy products inflation (ρ=-0.06). The inflation series with highest persistence are non-energy industrial goods and processed food, where ρ is respectively 0.44

and 0.40.[9] Administered prices and tobacco inflation (HICPADT) displays a very low level of persistence and under the time-varying approach it even displays negative correlation, thus contributing to reduction of the persistence of headline inflation. Another interesting observation is that HICPADT quarterly inflation is the only inflation series in Bulgaria with an overall declining trend over the 1998–2007 period (see Figure 7 below). This is an indication that the group of non-market determined prices is gradually becoming a factor with a decreasing impact on inflation in Bulgaria following the attainment of cost-recovery level for the major utilities comprising administered prices. Accordingly, all measures of core inflation exhibit higher persistence than headline inflation as the more volatile subcomponents of HICP reducing persistence (administered prices and tobacco, unprocessed food and energy products) are excluded from them. This finding is also confirmed by IPN conclusions which are discussed in more detail in Section 3.

Figure 7

HICP ADMINISTERED PRICES AND TOBACCO (ADT) INFLATION



[9]By statistical standards, however, non-energy industrial goods (HICPIG) and processed food (HICPFPR) inflation series are far from being highly persistent series because their persistence parameter is far below unity and is also below 0.75 (in which case doubts about a potential downward bias may arise).

3.3. Persistence estimates for year-on-year inflation rates at quarterly frequency

Persistence estimates for yearly inflation rates are presented in Table 4 and Table 4A in the Appendix. The results from both tests show that under the constant mean assumption, similarly to quarterly inflation rates, the least persistent are unprocessed food and energy series. According to the ADF test headline CPI/HICP and HICP excluding energy and unprocessed food inflation series are weakly persistent with ρ respectively 0.61 and 0.66. However, the PP test does not confirm this result. For all other series at both the aggregate and disaggregate level, the two unit root tests reveal a relatively high level of persistence because, as a general rule, ρ exceeds 0.7 and for many of the series the random walk hypothesis is not rejected. Non-energy industrial goods inflation and the core HICP excluding energy, food, administered prices and tobacco (HICPXEFADT) inflation have with strongest dependence on past inflation.

As we have seen from Figure 2 in the Appendix, a typical feature of most of the yearly inflation series are the clear changes in their mean level over the estimated period. This fact may be one of the important factors in the failure of ADF and PP tests to reject the null non-stationarity hypothesis.[10] Indeed, conducting unit root tests for the de-trended year-on-year inflation rates reduces persistence estimates for all series based on statistical evidence against the random walk hypothesis at the conventional significance levels (cf. Table 4 and Table 4A). However, in comparison with quarterly inflation rates, the downward impact of the use of time-varying mean levels is much less evident. This finding points to higher persistence of year-on-year growth rates *vis-à-vis* quarter-on-quarter ones.

As with persistence estimates for quarterly inflation, unprocessed food and energy appear to be the least persistent series. We also note that non-energy industrial goods inflation is the only de-trended annual inflation series for which weak statistical evidence (at the 10per cent level) against the non-stationarity hypothesis is given by the two tests. For all other series, the ADF and PP tests reject the hypothesis at lower levels, thus providing stronger evidence against high persistence.

Non-energy industrial goods inflation exhibits not only a well-defined upward trend when measured both as quarterly inflation and as annual inflation at quarterly basis, but it also shows the highest degree of persistence

[10]Another reason for the failure of the ADF test to reject the random walk hypothesis is the small power of this test in short samples and our sample comprises no more than 34–35 observations.

of all inflation series. At this stage of the analysis and with the currently employed econometric techniques, it is difficult to distinguish whether industrial goods inflation is indeed a random walk process or a stationary process with persistence close to 1. It may also be that the short time series available compounds these difficulties additionally. One certain conclusion, however, is that when forecasting year-on-year inflation rates in Bulgaria following the disaggregate approach, two important findings should be taken into consideration: the upward trending behavior present in some of the series over the past few years (except administered prices and tobacco, and energy, see Figure 2 in the Appendix) and the higher persistence of yearly inflation series *vis-à-vis* quarterly ones.

4. Inflation persistence in Bulgaria in the context of persistence in the euro area and the other NMS

The main conclusions of the IPN research project concerning macrolevel univariate studies of inflation persistence (based on quarterly inflation series) can be summarized as follows (Altissimo, Ehrmann and Smets, 2006):

- Inflation persistence in euro area countries can be described as moderate; a crucial finding is that in measuring persistence it is important to take into consideration statistically significant breaks in the mean of inflation (especially when the time period is relatively long), to allow for a time-varying mean or to estimate persistence over shorter sample periods characterized by stable monetary policy regimes.

- Core inflation (described as non-energy non-food inflation) is found to display a higher level of persistence than overall inflation.

- On account of methodological differences between CPI and HICP related mainly to the requirement for inclusion of sales prices in the latter index, the level of persistence in HICP inflation is generally lower than that measured according to national practices.

- Inflation persistence estimates vary substantially across countries and sectors. Differences across countries are attributed to various country-specific factors driving inflation and its persistence (these factors cannot be identified with univariate models). Heterogeneity across sectors is more pronounced than that across countries. Unprocessed food and energy products are by far the least persistent inflation series, whereas non-energy industrial goods and services are comparatively more persistent.

- The persistence of headline inflation is generally higher than the average of the persistence of its subcomponents. This is the so-called aggregation effect which takes place when the more persistent series have a relatively large weight in the overall consumer price index.

- A substantial amount of empirical evidence concerning persistence of administered prices inflation is still insufficient due to the lack of a common methodology in compiling a harmonized index of such prices until recently. The scarce evidence available shows that inflation series excluding the impact of administered prices generally show only a smaller reduction in persistence compared with headline series.

The results of our empirical analysis, presented in Section 2, are broadly consistent with the IPN evidence on inflation persistence in euro area countries. In our analysis we made use of the need to control for changes in the mean of inflation by allowing for a time-varying mean in each inflation series. Even without employing the time-varying approach, however, we obtained an overall low level of quarterly inflation persistence, which was further reduced when the persistence of the demeaned inflation series was estimated. The finding about moderate inflation persistence is also supported by a study which investigates inflation persistence in few of the NMS (Franta, Saxa and Šmídková, 2007). Even though that study focuses on structural rather than univariate autoregressive models, it also reaches the conclusion that inflation persistence in the NMS is broadly comparable with that in euro area countries once time-varying mean models are employed.

In our empirical analysis we also confirmed the IPN conclusions regarding the higher persistence of core inflation and the pronounced heterogeneity of inflation persistence across sectors, including the relatively higher persistence of services and industrial goods compared to that of unprocessed food and energy products.

We did not consider important to analyze the differences between the inflation inertia of the HICP and the national CPI indices in Bulgaria and we have focused only on the HICP for the following reasons. First, full harmonization with the HICP methodology was implemented only in 2005 and thus differences between CPI and HICP inflation series are substantial only over a relatively short time period. Second, one of the main objectives of our empirical evaluation of inflation persistence in Bulgaria was to identify whether this persistence might have an impact on the fulfillment of the Maastricht criterion on inflation which is based on the HICP and not on the national CPI. Therefore, inflation persistence of other variables analysed in IPN studies such as national CPI and GDP deflator has not been explored in this paper.

We did not make a thorough analysis of the impact of aggregation effects on inflation persistence in Bulgaria as the level of data we were working with was not sufficiently disaggregate. We note that some of the IPN studies have analyzed inflation persistence using most disaggregate price indices (e.g. Lünnemann and Mathä, 2004; Altissimo, Mojon, Zaffaroni, 2004).

Nevertheless, based on our current estimates we can tentatively conclude that headline inflation persistence is largely affected by the persistence of industrial goods and services inflation as they make up a substantial part of the consumer basket.

Concerning administered prices, our conclusion is somewhat different from the IPN finding that inflation excluding non-market prices has a lower persistence than overall inflation. All our results, including quarterly and yearly inflation rates, show that the persistence of HICP excluding administered prices and tobacco (HICPXADT) inflation is higher than that of headline CPI/HICP inflation. Two important factors, which differentiate Bulgaria from euro area countries, may account for this apparent contradiction of results. First, there have been more frequent and substantial adjustments in administered prices in Bulgaria. This makes the administered prices and tobacco (ADT) inflation series in Bulgaria more volatile, and hence, less persistent, than that of euro area countries. Second, the different results might reflect the higher weight of administered prices in Bulgarian overall index which makes the impact of their exclusion on headline inflation persistence more pronounced.

To sum up, the empirical analysis of inflation persistence conducted for Bulgaria supports the main findings of the IPN network. Unfortunately, it is not straightforward to compare the degree of persistence of our inflation directly with the already obtained estimates for the euro area. Although such comparison is undoubtedly useful, it is obstructed by various differences such as estimation techniques and time periods covered. For example, as we have already noted, most IPN studies obtain their main persistence estimates when controlling for structural shifts in the mean level of inflation. This should have been an important consideration in these studies, accounting for the fact that estimation has been done over long time periods encompassing changes in the monetary policy regime. Obviously, such a consideration is less relevant for the period we estimate for Bulgaria which has not been characterized by one or two discrete shifts in the mean of inflation. Inflation dynamics has been driven not by one-off factors but by different overlapping processes such as the macroeconomic stabilization following the introduction of the currency board, the various and continuous adjustments in administered prices, the catching-up growth, etc. For this reason, in measuring persistence we have opted for controlling for the different trending pattern of each inflation variable. Consequently, we have not applied an OLS estimation which can flexibly handle introduction of various shift dummy variables. Our approach has been to use the automatic ADF and PP tests both for the constant mean estimations and for the time-varying ones (i.e. for the demeaned series). In addition, while some of the IPN studies have controlled for the seasonal variation in the data again by introducing dummy variables,

in our case we have applied seasonal adjustment procedures before running the unit root tests.

Based on the above arguments, we do not compare our results directly with any of the IPN ones (for a detailed overview of the latter, the reader can refer to Altissimo, Ehrmann and Smets, 2006, pp. 13–20). We decide to make additional estimates on inflation persistence in euro area countries and the other new EU member states based on the methodological approach we have applied for Bulgaria. More specifically, we estimate inflation persistence for annualized quarterly HICP inflation rates for all countries over the 1998–2007 period. The inflation series for all countries are deseasonalized with Census X-12 procedure, with additive adjustment. Then, with a view to adopting a unified approach regarding the mean of the inflation series for each country, we make the simplified assumption that inflation rates for all countries have had a time-varying mean and estimate equations (6) and (7) under this assumption. We obtain the time-varying mean of inflation for each country again by employing the HP filter to the deseasonalized series (the smoothing parameter is set to 1600). Persistence results, as measured by the parameter ρ obtained indirectly from the OLS estimation of the PP test (equation 7) are presented in Table 5.[11]

Table 5. Estimates of inflation peristence /time-varying mean/[1]
Annualized quarter-on-quarter HICP inflation rates, seasonally adjusted
Sample period: 1998:q1 - 2007:q4

Equation (7) /PP unit root test, no intercept included/

euro area	ρ	*New Member States*	ρ
Belgium	**-0.27**	Bulgaria	**0.13**
Germany	**-0.15**	The Czech Republic	**0.34**
Ireland	**0.17**	Estonia	**0.22**
Greece	**-0.04**	Hungary	**0.30**
Spain	**-0.15**	Latvia	**0.51**
France	**-0.26**	Lithuania	**0.22**
Italy	**-0.36**	Poland	**0.50**
Luxemburg	**-0.13**	Romania	**0.37**
The Netherlands	**0.09**	Slovakia	**-0.03**
Austria	**0.08**		
Portugal	**0.11**		
Finland	**0.08**		
Slovenia	**0.06**		
Malta	**-0.02**		
Cyprus	**0.01**		

[11]Persistence estimates produced by the ADF unit root test are not qualitatively different from those obtained by the PP test as the automatic selection of lag length according to BIC is most often zero. In other words, in the ADF test we have had the case of AR(1) inflation processes.

It can be seen that inflation persistence in euro area countries is much lower than that observed in the NMS. One factor which may explain these differences is the monetary policy conducted by the ECB which anchors inflation expectations, thus reducing overall dependence on past inflation developments (Altissimo, Ehrmann and Smets, 2006). The variation in inflation persistence across euro area countries, however, points to the importance of other factors underlying inflation inertia as well (e.g. we have a difference of 0.53 between the lowest serial correlation in inflation observed for Italy, -0.36, and the highest, 0.17, respectively for Ireland).

It is noteworthy that the three most recent euro area members have had almost no persistence in their inflation rates over the estimated period. Similarly, Slovakia, the country which stands closest to meeting the Maastricht inflation requirement, has the lowest persistence estimate from all NMS (-0.03). Bulgaria is the second county within the NMS group with lowest inertia in inflation (0.13). This estimate for Bulgaria is below the estimates of persistence for Estonia (0.22), Lithuania (0.22), the Czech Republic (0.34), Romania (0.37), Poland (0.50) and Latvia (0.51). The univariate models we have employed to produce the results on inflation persistence do not allow us to disentangle the various sources of these differences in persistence. One can argue that the observed heterogeneity in the degree of inflation inertia across the NMS to a certain extent reflects the differences in the levels and variability of their inflation rates over the period. Other related factors that need to be considered are the differences in monetary policy regimes across the countries. We also note that for the Czech Republic where inflation targeting has been pursued over a relatively long period of time and the criterion on inflation has generally been met, a persistence estimate of 0.34 might have a different connotation from the estimate of 0.37 in Romania which has experienced longer periods of higher inflation.[12] Some of the variation in the persistence estimates across the NMS may also be due to certain differences in their HICP harmonization process which might have affected the methodological characteristics of their HICP measurement and respectively their inflation persistence. While we are unable to make a conclusive explanation of the differences in inflation persistence in the NMS, we can still conclude that Bulgaria is in a relatively good position in terms of inflation persistence both within the NMS group and the euro area countries.

[12]Inflation targeting has been the monetary policy regime in Romania too, though only over the last couple of years.

5. Analysis of price dynamics in Bulgaria

In this section we analyze price dynamics in Bulgaria using monthly data of representative goods prices given as averages for the whole country over the period 2000-2007. This data is further aggregated into the main HICP groupings under the Eurostat classification. The observed patterns of price adjustments are discussed in the context of the IPN micro evidence on the frequency and size of individual price changes in euro area countries over the period 1998-2004. It is worthwhile noting that the differences in the nature of the data (averages of individual prices in Bulgaria vs. individual prices in the euro area) do not allow making definite conclusions and only provide general idea of the respective price dynamics processes.

For Bulgaria, we use the monthly data of representative goods prices collected by the Bulgarian National Statistical Institute (NSI) to compute the Bulgarian Consumer Price Index (CPI) and the Harmonized Index of Consumer Prices (HICP). These are averages for the whole country of actually observed individual prices by the surveyors when they visit stores. For each month in the period 2000 – 2007 the NSI dataset comprises between 512 and 548 average prices of representative goods and services. We round the average prices of these products to hundredths to eliminate those changes which are less then one hundredth of the Bulgarian Lev.

The representative goods and services are divided in 3 main groups: food (processed and unprocessed), non-food industrial goods (energy and non-energy industrial goods) and services (catering services and others).

Then we calculate the share of prices which are changed each month as the number of product price changes (for a given product group or the overall consumer basket) divided by the total number of product prices. Below, we use the terms 'share of changed prices' and 'frequency of price changes' interchangeably. In addition, we also study the average size of price changes (respectively, the average size of price increases and of price decreases) observed in the data over the period under investigation.

The frequency and size of consumer price changes in Bulgarian data are then compared with the respective results obtained for the euro area. For the latter, we use the data summarized by Altissimo, Ehrmann and Smets (2006).[13]

Table 6 below presents an overview of the main characteristics of consumer price changes in Bulgaria and euro area countries. Table 7 provides additional information on the frequency of consumer price changes by product type for different countries in the euro area.

[13]For comparability of our results with those at euro area level, we do not exclude administered prices and tobacco from the total group of representative goods and services observed by the National Statistical Institute of Bulgaria.

Table 6

MAIN CHARACTERISTICS OF CONSUMER PRICE CHANGES IN BULGARIA AND THE EURO AREA (AVERAGE PER MONTH, %)

	Share of goods with increased prices	Share of goods with decreases prices	Share of goods with changed prices	Average size of price increases	Average size of price decreased	Average size of price increased (euro area)	Average size of price decreases (euro area)
Food	43	27	69	6	-6		
processed	35	18	53	2	-2	7	-8
unprocessed	52	38	90	8	-7	15	-16
Non-energy industrial goods	40	34	74	2	-1	9	-11
Energy	54	37	91	3	-2	3	-2
Services	35	12	48	6	-3	7	-9
services excl. catering	32	11	43	8	-4		
catering	45	16	61	2	-1		
Overall CPI/HICP	**40**	**26**	**66**	**4**	**-3**		
Euro area	***9***	***6***	***15***	***8***	***-10***		

Notes: The data for Bulgaria are based on monthly prices at representative products level. The data for the euro area are based on monthly individual price records (a more disaggregated level than the representative products level).

For Bulgaria, the table reports averages for the 2000-2007 period, while for the euro area the results refer to averages for the 1998-2004 period.

Unsurprisingly, prices in Bulgaria change much more frequently than in euro area countries. On average over the period, 66per cent of all prices (at representative products level) are changed each month in Bulgaria. In the euro area, the corresponding frequency of price adjustments is equal to 15per cent. The higher frequency of price changes in Bulgaria vis-?-vis the euro area is largely accounted for by the processes of catching-up growth and price level convergence the country has been undergoing.[14]

[14]Additional sources of price changes in Bulgaria, which differentiate the country from the euro area, include, for example, the much more frequent adjustments in administered prices.

Notwithstanding these differences, the observed pattern of price changes across product groups in Bulgaria is the same as in euro area countries. As shown in Tables 6 and 7, fuels and unprocessed food are the product groups with the most frequently changed prices each month both in Bulgaria and the euro area. On average over the period, 91per cent of all energy prices and 90per cent of all unprocessed food prices are changed each month in Bulgaria, whereas in the euro area the corresponding shares are 78per cent and 28per cent respectively. The main reason for the higher frequency of changes in energy and unprocessed food prices in comparison with the other product groups is their quick response to changes in input prices. It can also be seen that services prices represent by far the least frequently changed group of prices, with frequency of price changes of 48per cent in Bulgaria and 5.6per cent in the euro area. Not unexpectedly, in Bulgaria industrial goods are undergoing far more price changes than in the euro area where this group of products is characterized by much greater rigidities.

There is a substantial degree of heterogeneity across sectors between the shares of price increases and price decreases in all price changes. In Bulgaria price increases are much more common than price decreases for the services group and this is also typical for euro area countries (see Table 8). Decreases of prices in services are less then in the other groups due to their non-tradable nature, higher labour input share and possible downward nominal wage rigidity.[15] On the whole, in terms of the share of price increases in all price changes, similarities across product groups are much stronger than differences. In this respect, product characteristics appear more pronounced than country group differences.

The euro area and Bulgaria differ in the average size of price increases and price decreases. Surprisingly, a noticeable feature of price changes in Bulgaria is that the magnitude of these changes is smaller than that in the euro area (see Table 6). Except for energy products, the size of price decreases is much smaller for most product groups in Bulgaria. We note, however, that conclusive statements about the size of price changes cannot not be made because of the different nature of the data (averages of individual prices in Bulgaria vs. individual prices themselves in the euro area).

[15]For explanation of the factors underlying the lower frequency of price decreases in the services sector, see Altissimo, Ehrmann and Smets (2006).

Table 7

FREQUENCY OF CONSUMER PRICE CHANGES BY PRODUCT TYPE (AVERAGE PER MONTH,%)

	Unprocessed food	Processed food	Fuels goods	Industrial	Services	Total
Belgium	31.5	19.1	81.6	5.9	3.0	17.6
Germany	25.2	8.9	91.4	5.4	4.3	13.5
Spain	50.9	17.7	n.a.	6.1	4.6	13.3
France	24.7	20.3	76.9	18.0	7.4	20.9
Italy	19.3	9.4	61.6	5.8	4.6	10.0
Luxembourg	54.6	10.5	73.9	14.5	4.8	23.0
The Netherlands	30.8	17.3	72.6	14.2	7.9	16.2
Austria	37.5	15.5	72.3	8.4	7.1	15.4
Portugal	55.3	24.5	15.9	14.3	13.6	21.1
Finland	52.7	12.8	89.3	18.1	11.6	20.3
Euro area	**28.3**	**13.7**	**78.0**	**9.2**	**5.6**	**15.1**
Bulgaria	89.5	53.2	91.1	73.9	47.9	65.6

Table 8

SHARE OF PRICE INCREASES IN ALL PRICE CHANGES (AVERAGE PER MONTH,%)

	Unprocessed food	Processed food	Fuels goods	Industrial	Services	Total
Bulgaria	58	67	59	54	74	61
Euro area	**54**	**54**	**54**	**57**	**80**	**58**

In Figures 8 through 11 below we present price dynamics by product type in Bulgaria based on monthly data over the period from 2000 to 2007. In general, there is an overall tendency to fewer price changes or greater price stability until 2005. Since then, the share of changed prices (respectively the share of price increases) has been increasing. Price reductions have not been uncommon; however, since 2004 their share in overall price changes has been decreasing considerably. The increasing share of upward price adjustments since 2005 explains the higher inflation rate over the last years and can be clearly seen in the figures with price increases and price decreases. This tendency also helps explain the upward trending pattern in the inflation rates of some of the CPI/HICP components which was documented in Section 2.

Figure 12 displays that indeed there is some positive correlation between the overall CPI/HICP monthly inflation rate and the proportion of price increases relative to price decreases over the period 2000–2007. An alternative explanation of the link between inflation and the frequency of price adjustments is offered by Vermeulen et.al. (2007) who suggest that such a link might indicate the existence of state-dependent pricing or the impact of aggregate economic shocks (in this case, inflation) on the pricing behaviour of firms. We believe that because of the nature of price data for Bulgaria (at representative products level and not at individual price records level), the validity of such conclusion for our country needs to be confirmed by further analysis supported by more disaggregate data.

As it is confirmed for other EU countries, monthly frequencies of price changes exhibit clear seasonal patterns. As a whole, price changes are more likely to take place during the first quarter (especially in January) or after the summer period (especially in September) and are less frequent in the summer (see Table 9 in the Appendix).

Figure 8

OVERALL PRICE DYNAMICS, CPI/HICP

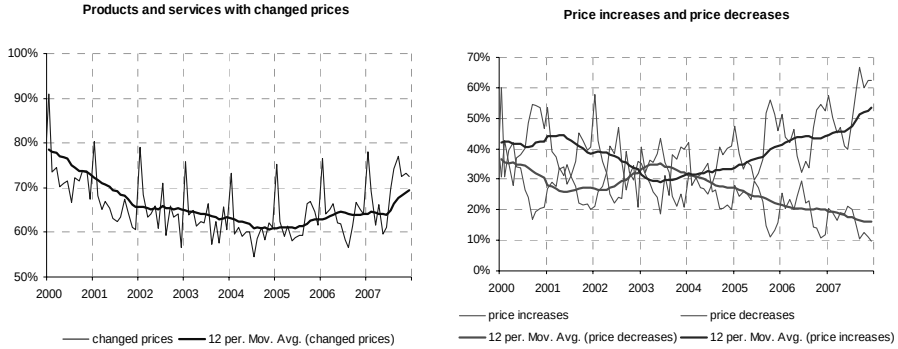
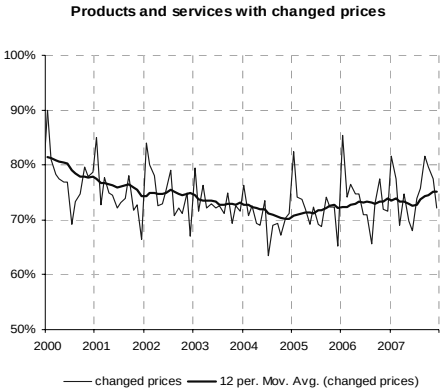


Figure 9

PRICE DYNAMICS OF NON-ENERGY INDUSTRIAL GOODS



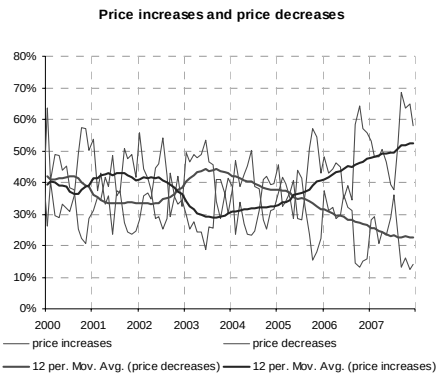
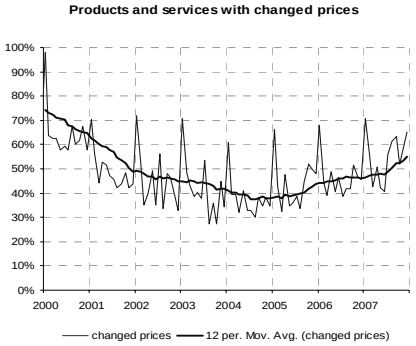


Figure 10

PRICE DYNAMICS OF SERVICES



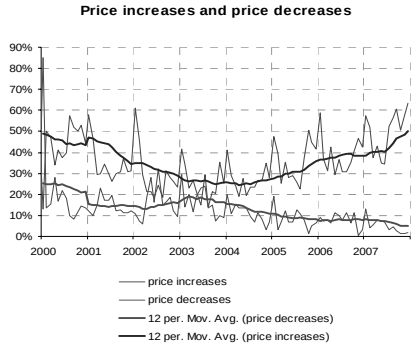


Figure 11

PRICE DYNAMICS OF FOOD PRODUCTS

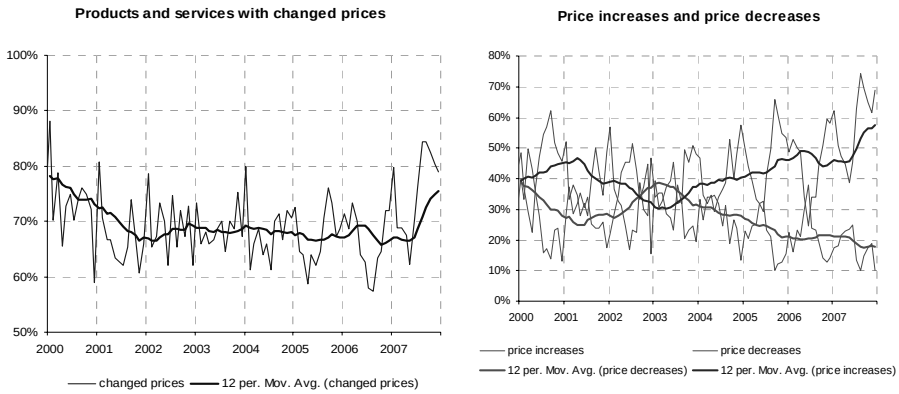
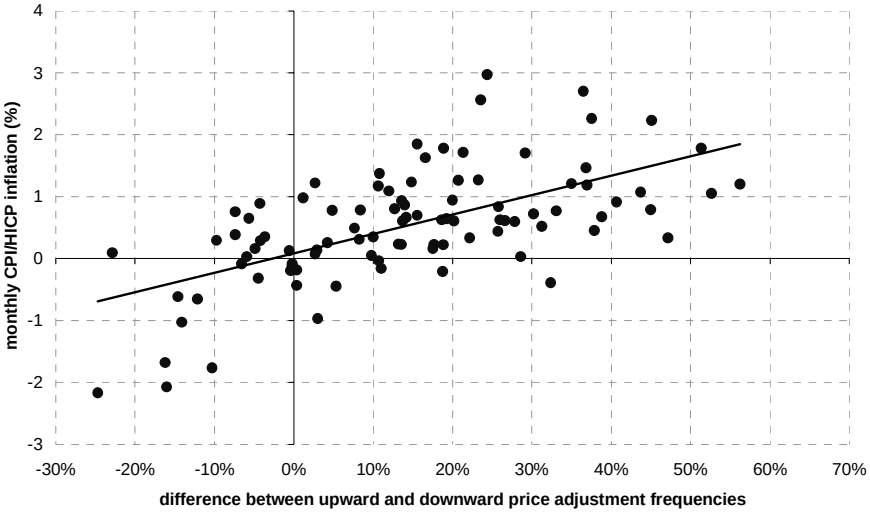


Figure 12

MONTHLY CPI/HICP INFLATION AND THE DIFFERENCE BETWEEN UPWARD AND DOWNWARD PRICE ADJUSTMENTS PER MONTH, FEBRUARY2000 – DECEMBER 2007



6. Conclusions

This paper has investigated empirically inflation persistence in Bulgaria following the two main directions of research undertaken under the Inflation Persistence Network (IPN) project of the Eurosystem: univariate models of inflation and analysis of price dynamics based on disaggregate data over the 2000-2007 period.

Persistence estimates were based on the sum of autoregressive coefficients in time-series equations of inflation. They were derived indirectly from the Augmented Dickey-Fuller test and the Phillips-Perron unit root tests. The variables used in the empirical analysis included headline HICP inflation, three measures of core inflation and inflation of the main HICP sub-components according to the Eurostat classification. Inflation rates were measured over the period from 1998 to 2007 as quarter-on-quarter inflation (non-seasonally adjusted and seasonally adjusted) and as year-on-year rate of change at quarterly frequency.

Preliminary persistence estimates were obtained under the assumption of a constant mean of inflation. The underlying assumption for the main results on persistence was that of a time-varying mean which was considered more appropriate for Bulgarian inflation series than the standard for euro area countries approach of controlling for structural break(s) in inflation. This assumption was motivated by the fact that as a country undergoing a process of economic transformation, catching-up growth and real income and price convergence, Bulgaria has been characterized by trending behavior of most inflation rates. Persistence under the time-varying approach was calculated again with the ADF and the PP tests which this time were applied to de-trended inflation series. Trends in inflation rates were proxied by HP filter-based trend components.

Our main results on inflation persistence in Bulgaria can be summarized as follows. Bulgaria is characterized by an overall low level of persistence of quarterly inflation. Even under the na?ve (constant mean) assumption, at the aggregate level persistence estimates for both non seasonally-adjusted and seasonally-adjusted inflation series do not exceeded 0.5. This provides not only strong evidence against high persistence but also against the problem of downward bias of the estimates. At the disaggregate level, the constant mean assumption produces mixed results, with unprocessed food and energy standing out as least persistent inflation series and industrial goods as showing relatively higher persistence than the other series.

We have found that the quarterly inflation series of industrial goods, services and a core index, including only these two subcomponents, show a clear-cut upward trending pattern over the last three years which contributes

to considerable overestimation of their persistence under the constant mean approach. When persistence is measured for de-trended series, persistence estimates are significantly reduced. Deseasonalized inflation series are more persistent than the non-seasonally adjusted ones because of the smaller variability observed in them once seasonal variation is removed.

Our final results for deseasonalized and detrended quarterly inflation series reveal a low degree of inflation persistence not exceeding 0.44 across-the-board. At the aggregate level, headline inflation is characterized by very low persistence (ρ=0.13) mainly due to the negative serial correlation evident for unprocessed food and energy products. The inflation series with highest persistence are industrial goods (ρ=0.44) and processed food (0.40). All three core inflation variables exhibit more inertia than headline inflation.

Even though de-trended year-on-year inflation rates show relatively higher persistence than quarterly series, statistical evidence at conventional significance levels supports the lack of strong persistence in these series. Among all inflation series at the aggregate and sectoral level, industrial goods display strongest dependence on past inflation. Calculation of unbiased estimates for the persistence of yearly inflation series, which could shed more light on the exact degree of persistence in these series, is a promising area of further empirical research.

The empirical analysis of quarterly inflation persistence in Bulgaria as a whole supports the main findings of the IPN for euro area countries. The only contradiction with the IPN is that persistence of inflation excluding administered prices and tobacco is higher than overall inflation, whereas the IPN evidence has the opposite result. We attribute this discrepancy in results to the higher volatility (hence, lower persistence) of non-market prices inflation in Bulgaria and the heavier weight of this group in the consumer basket. This leads to higher persistence of core inflation excluding administered prices.

We also make direct comparison with persistence estimates for euro are countries and the other New Member States (NMS) following the same methodological approach adopted for Bulgaria (i.e. application of unit root tests to seasonally adjusted and de-trended quarterly inflation series over the 1998-2007 period). The results show that Bulgaria has the second smallest persistence from all NMS following Slovakia. In addition, it is also in a relatively good position compared with some of euro area countries.

Not unexpectedly, the analysis of price dynamics shows that prices in Bulgaria change much more frequently than in euro area countries. On average over the 2000–2007 period, 66per cent of all prices at representative products level change each month in Bulgaria, whereas in the euro area only

15per cent of all prices based on individual price records level are changed. The higher frequency of price changes in Bulgaria *vis-à-vis* the euro area is largely accounted for by the processes of catching-up growth and price level convergence the country has been undergoing. Despite these differences in the frequencies of price adjustments, the observed pattern across product groups is the same as in euro area countries. For example, energy products and unprocessed food have the highest frequency of price changes, and services - the lowest frequency, respectively.

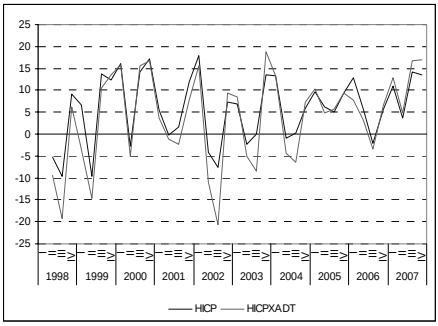
In terms of symmetry between the shares of price increases and price decreases in all price changes, on the whole, similarities across product groups are again much more pronounced than differences across countries. The share of price increases is much higher in the services group compared with the other product groups both in Bulgaria and in the euro area. This could be accounted for by the non-tradable nature of services, the higher labour input share in these products and the possible downward nominal wage rigidities in the sector.

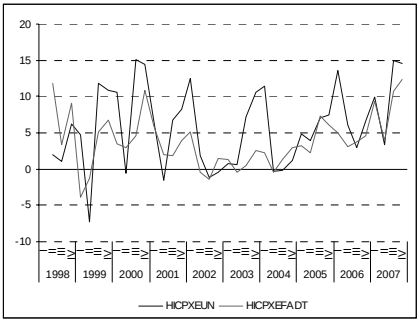
The analysis of price dynamics in Bulgaria based on monthly data over the 2000–2007 period reveals a tendency to fewer price changes (i.e. greater price stability) from 2000 until 2004 which is then followed by an increasing trend of price increases relative to downward price adjustments. This has been common for most of the CPI/HICP components and helps explain the acceleration of core inflation since 2005. Further analysis could focus on the price-setting behaviour of Bulgarian firms and more particularly on investigating the extent to which inflation as a macroeconomic shock affects the frequency of price changes and the respective distribution between price increases and price decreases. Such an analysis will provide evidence for the relative importance of state-dependent time-dependent pricing and will shed more light on inflation processes in the country.

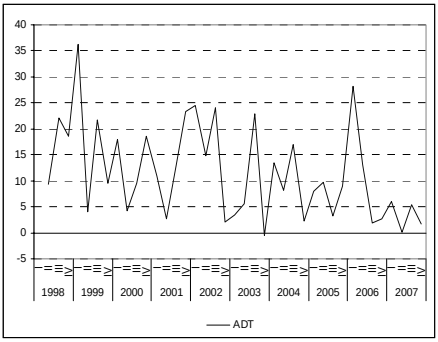
APPENDICES

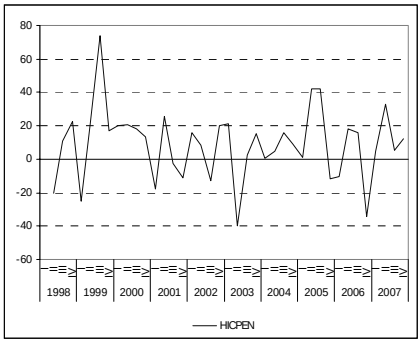
Figure1

ANNUALIZED QUARTER-ON-QUARTER INFLATION RATES[16]



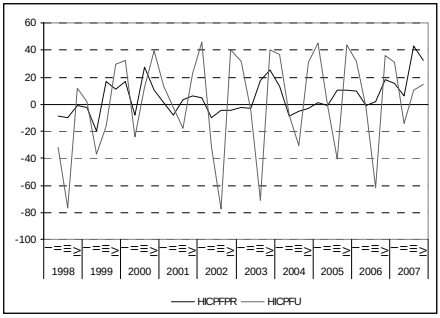


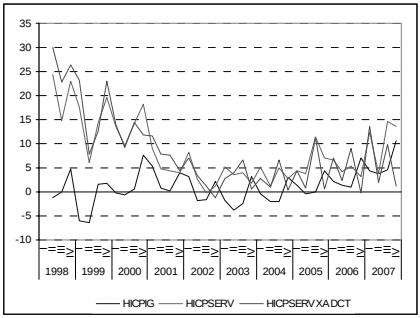




CPI/HICP	Harmonized index of consumer prices (HICP), HICP is equal to the national CPI until 2004
HICPXADT	HICP excluding administered prices and tobacco
HICPXEUN	HICP excluding energy and unprocessed food
HICPXEFADT	HICP excluding energy, food, administered prices and tobacco
ADT	HICP administered prices and tobacco
HICPEN	HICP energy excluding administered prices

[16]Inflation rates are not seasonally-adjusted. The following standard transformation is applied for obtaining inflation rates: ?t = 400*ln(Pt/ Pt-1), where ?t is the annualized quarter-on-quarter inflation rate and Pt is the respective consumer price index at the aggregate and disaggregate (sectoral) level, 2005=100.

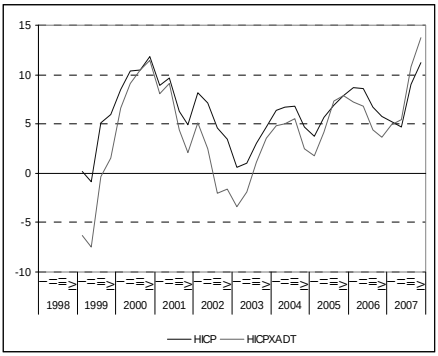


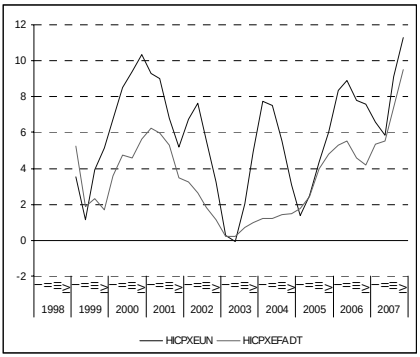


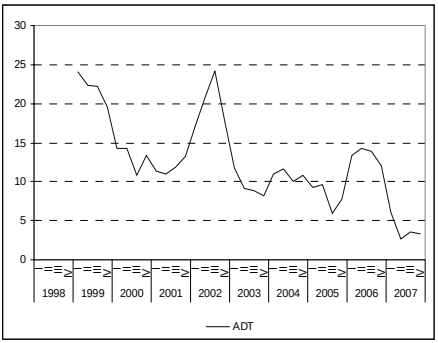
HICPFPR	HICP processed food excluding tobacco
HICPFU	HICP unprocessed food
HICPIG	HICP non-energy industrial goods excluding administered prices
HICPSERV	HICP services excluding administered prices
HICPSERVXADCT	HICP services excluding administered prices, catering and transport services

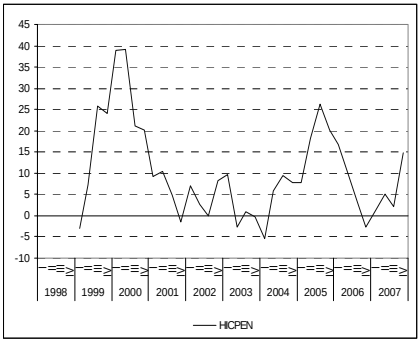
Figure 2

YEAR-ON-YEAR INFLATION RATES /AT QUARTERLY FREQUENCY/

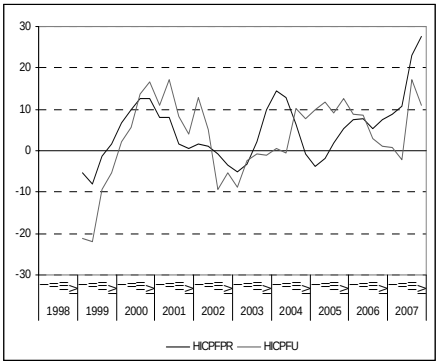


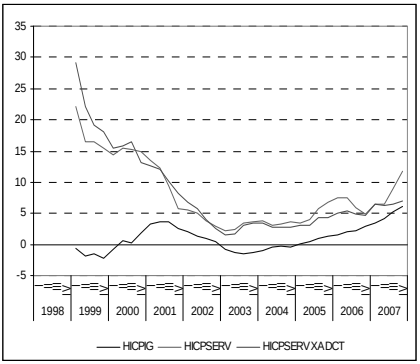






CPI/HICP	Harmonized index of consumer prices (HICP), HICP is equal to the national CPI until 2004
HICPXADT	HICP excluding administered prices and tobacco
HICPXEUN	HICP excluding energy and unprocessed food
HICPXEFADT	HICP excluding energy, food, administered prices and tobacco
ADT	HICP administered prices and tobacco
HICPEN	HICP energy excluding administered prices





HICPFPR	HICP processed food excluding tobacco
HICPFU	HICP unprocessed food
HICPIG	HICP non-energy industrial goods excluding administered prices
HICPSERV	HICP services excluding administered prices
HICPSERVXADCT	HICP services excluding administered prices, catering and transport services

Table 1. Estimates of inflation peristence /constant mean/
Annualized quarter-on-quarter inflation rates, non-seasonally adjusted
Sample period: 1998:q1 - 2007:q4

Equation (2) /ADF unit root test, intercept included/

Inflation variable	ρ-1	ρ	Lags	P-value	t-statistics
aggregate level					
CPI/HICP	-1.38	**-0.38**	1	0.00	-7.09***
HICPXADT	-1.21	**-0.21**	1	0.00	-6.28***
HICPXEUN	-0.78	**0.22**	0	0.00	-4.68***
HICPXEFADT	-0.65	**0.35**	0	0.00	-4.12***
ADT	-0.93	**0.07**	0	0.00	-5.53***
disaggregate level					
HICPEN	-1.38	**-0.38**	1	0.00	-6.18***
HICPFPR	-0.62	**0.38**	0	0.01	-3.82***
HICPFU	-1.20	**-0.20**	3	0.03	-3.13**
HICPIG	-0.27	0.73	5	0.70	-1.11
HICPSERV	-0.45	**0.55**	0	0.01	-3.66***
HICPSERVXADCT	-0.57	**0.43**	3	0.00	-4.86***

Equation (5) /PP unit root test, intercept included/

Inflation variable	ρ-1	ρ	Lags	P-value	t-statistics
aggregate level					
CPI/HICP	-0.90	**0.10**	0	0.00	-6.70***
HICPXADT	-0.80	**0.20**	0	0.00	-4.82***
HICPXEUN	-0.78	**0.22**	0	0.00	-4.50***
HICPXEFADT	-0.65	**0.35**	0	0.00	-4.13***
ADT	-0.93	**0.07**	0	0.00	-5.53***
disaggregate level					
HICPEN	-1.01	**-0.01**	0	0.00	-6.31***
HICPFPR	-0.62	**0.38**	0	0.01	-3.72***
HICPFU	-1.01	**-0.01**	0	0.00	-7.90***
HICPIG	-0.61	**0.39**	0	0.02	-3.34**
HICPSERV	-0.45	**0.55**	0	0.01	-3.66***
HICPSERVXADCT	-0.45	**0.55**	0	0.01	-3.75***

Notes:

CPI/HICP	Harmonized index of consumer prices (HICP), HICP is equal to the national CPI until 2004
HICPXADT	HICP excluding administered prices and tobacco
HICPXEUN	HICP excluding energy and unprocessed food
HICPXEFADT	HICP excluding energy, food, administered prices and tobacco
ADT	HICP administered prices and tobacco
HICPEN	HICP energy, excluding administered prices
HICPFPR	HICP processed food, excluding tobacco
HICPFU	HICP unprocessed food
HICPIG	HICP non-energy industrial goods, excluding administered prices
HICPSERV	HICP services, excluding administered prices
HICPSERVXADCT	HICP services, excluding administered prices, catering and transport services

1) (ρ-1) is the coefficient estimated by the ADF and PP tests based respectively on equations (2) and (5)
2) ρ is the sum of autoregressive coefficients derived indirectly from (ρ-1)
3) The lag length in the ADF test is selected automatically by the Schwarz Info Criterion (BIC), max length=5
4) The default options of EViews are used for the spectral estimation method and the bandwidth in the PP test
5) P-value is the probability associated with the null hypothesis of a unit root in the respective inflation series
6) t-statistics is the statistics of the coefficient (ρ-1) in the OLS equation estimated by the tests
7) MacKinnon's critical values are used for the result on the t-statistics
8) ***, ** and * denote rejection of the non-stationarity null at the 1%, 5% and 10% respectively
9) Numbers in bold indicate that it is possible to reject statistically that ρ=1

Table 1A. Estimates of inflation peristence /time-varying mean/[1]
Annualized quarter-on-quarter inflation rates, non-seasonally adjusted
Sample period: 1998:q1 - 2007:q4

Equation (6) /ADF unit root test, no intercept included/

Inflation variable	ρ-1	ρ	Lags	P-value	t-statistics
aggregate level					
CPI/HICP	-1.47	**-0.47**	1	0.00	-7.46***
HICPXADT	-1.37	**-0.37**	1	0.00	-7.10***
HICPXEUN	-1.11	**-0.11**	1	0.00	-5.04***
HICPXEFADT	-0.90	**0.10**	0	0.00	-5.67***
ADT	-1.14	**-0.14**	0	0.00	-7.10***
disaggregate level					
HICPEN	-1.47	**-0.47**	1	0.00	-6.71***
HICPFPR	-0.82	**0.18**	0	0.00	-5.02***
HICPFU	-1.12	**-0.12**	3	0.00	-2.48**
HICPIG	-0.65	**0.35**	5	0.02	-2.39**
HICPSERV	-1.02	**-0.02**	0	0.00	-6.27***
HICPSERVXADCT	-1.30	**-0.30**	3	0.00	-3.28***

Equation (7) /PP unit root test, no intercept included/

Inflation variable	ρ-1	ρ	Lags	P-value	t-statistics
aggregate level					
CPI/HICP	-0.96	**0.04**	0	0.00	-8.43***
HICPXADT	-0.91	**0.09**	0	0.00	-6.00***
HICPXEUN	-0.87	**0.13**	0	0.00	-5.92***
HICPXEFADT	-0.90	**0.10**	0	0.00	5.67***
ADT	-1.14	**-0.14**	0	0.00	-7.63***
disaggregate level					
HICPEN	-1.04	**-0.04**	0	0.00	-6.53***
HICPFPR	-0.82	**0.18**	0	0.00	-4.92***
HICPFU	-1.06	**-0.06**	0	0.00	-8.75***
HICPIG	-0.84	**0.16**	0	0.00	-4.87***
HICPSERV	-1.02	**-0.02**	0	0.00	-6.30***
HICPSERVXADCT	-1.27	**-0.27**	0	0.00	-8.19***

Notes:
Inflation variables are denoted as explained in Table 1.

1) (ρ-1) is the coefficient estimated by the ADF and PP tests based respectively on equations (6) and (7)
2) ρ is the sum of autoregressive coefficients derived indirectly from (ρ-1)
3) The lag length in the ADF test is selected automatically by the Schwarz Info Criterion (BIC), max length=5
4) The default options of EViews are used for the spectral estimation method and the bandwidth in the PP test
5) P-value is the probability associated with the null hypothesis of a unit root in the respective inflation series
6) t-statistics is the statistics of the coefficient (ρ-1) in the OLS equation estimated by the tests
7) MacKinnon's critical values are used for the result on the t-statistics
8) ***, ** and * denote rejection of the non-stationarity null at the 1%, 5% and 10% respectively
9) Numbers in bold indicate that it is possible to reject statistically that ρ=1

[1]ADF and PP unit root tests are applied to the deviation of inflation from its time-varying mean
The time-varying mean of an inflation series is proxied by HP filter-based trend component
The smoothing parameter is set to 1600, the standard number for quarterly data

Table 2. Estimates of inflation peristence /constant mean/
Annualized quarter-on-quarter inflation rates, seasonally adjusted [1]
Sample period: 1998:q1 - 2007:q4

Equation (2) /ADF unit root test, intercept included/

Inflation variable	ρ-1	ρ	Lags	P-value	t-statistics
aggregate level					
CPI/HICP	-0.76	**0.24**	0	0.00	-4.62***
HICPXADT	-0.64	**0.36**	0	0.00	-4.02***
HICPXEUN	-0.48	**0.52**	0	0.03	-3.17**
HICPXEFADT	-0.53	**0.47**	0	0.00	-3.85***
ADT	-	-	-	-	-
disaggregate level					
HICPEN	-1.03	**-0.03**	0	0.00	-6.40***
HICPFPR	-0.38	**0.62**	0	0.09	-2.68*
HICPFU	-0.99	**0.01**	0	0.00	-5.85***
HICPIG	-0.26	0.74	1	0.48	-1.59
HICPSERV	-0.34	**0.66**	0	0.01	-3.70***
HICPSERVXADCT	-0.33	**0.67**	2	0.00	-4.19***

Equation (5) /PP unit root test, intercept included/

Inflation variable	ρ-1	ρ	Lags	P-value	t-statistics
aggregate level					
CPI/HICP	-0.76	**0.24**	0	0.00	-4.62***
HICPXADT	-0.64	**0.36**	0	0.00	-4.06***
HICPXEUN	-0.48	**0.52**	0	0.02	-3.32**
HICPXEFADT	-0.53	**0.47**	0	0.00	-4.17***
ADT	-	-	-	-	-
disaggregate level					
HICPEN	-1.03	**-0.03**	0	0.00	-6.44***
HICPFPR	-0.38	**0.62**	0	0.08	-2.73*
HICPFU	-0.99	**0.01**	0	0.00	-5.88***
HICPIG	-0.29	0.71	0	0.27	-2.04
HICPSERV	-0.34	**0.66**	0	0.01	-3.70***
HICPSERVXADCT	-0.27	**0.73**	0	0.00	-4.43***

Notes:
Inflation variables are denoted as explained in Table 1.

1) (ρ-1) is the coefficient estimated by the ADF and PP tests based respectively on equations (2) and (5)
2) ρ is the sum of autoregressive coefficients derived indirectly from (ρ-1)
3) The lag length in the ADF test is selected automatically by the Schwarz Info Criterion (BIC), max length=5
4) The default options of EViews are used for the spectral estimation method and the bandwidth in the PP test
5) P-value is the probability associated with the null hypothesis of a unit root in the respective inflation series
6) t-statistics is the statistics of the coefficient (ρ-1) in the OLS equation estimated by the tests
7) MacKinnon's critical values are used for the result on the t-statistics
8) ***, ** and * denote rejection of the non-stationarity null at the 1%, 5% and 10% respectively
9) Numbers in bold indicate that it is possible to reject statistically that ρ=1

[1]Inflation series are seasonally-adjusted with Census-X12 procedure, using additive adjustment
ADT (administered prices and tobacco) inflation is not seasonally adjusted
No estimates of inflation persistence are made for ADT

Table 2A. Estimates of inflation peristence /time-varying mean/[1]
Annualized quarter-on-quarter inflation rates, seasonally adjusted
Sample period: 1998:q1 - 2007:q4

Equation (6) /ADF unit root test, no intercept included/

Inflation variable	ρ-1	ρ	Lags	P-value	t-statistics
aggregate level					
CPI/HICP	-0.87	**0.13**	0	0.00	-5.33***
HICPXADT	-0.82	**0.18**	0	0.00	-5.06***
HICPXEUN	-0.60	**0.40**	0	0.00	-3.95***
HICPXEFADT	-0.79	**0.21**	0	0.00	-5.52***
ADT	-	-	-	-	-
disaggregate level					
HICPEN	-1.06	**-0.06**	0	0.00	-6.70***
HICPFPR	-0.60	**0.40**	0	0.00	-3.93***
HICPFU	-1.17	**-0.17**	0	0.00	-7.02***
HICPIG	-0.56	**0.44**	0	0.00	-3.75***
HICPSERV	-0.85	**0.15**	0	0.00	-5.69***
HICPSERVXADCT	-0.86	**0.14**	0	0.00	-6.04***

Equation (7) /PP unit root test, no intercept included/

Inflation variable	ρ-1	ρ	Lags	P-value	t-statistics
aggregate level					
CPI/HICP	-0.87	**0.13**	0	0.00	-5.34***
HICPXADT	-0.82	**0.18**	0	0.00	-5.14***
HICPXEUN	-0.60	**0.40**	0	0.00	-4.02***
HICPXEFADT	-0.79	**0.21**	0	0.00	-5.55***
ADT	-	-	-	-	-
disaggregate level					
HICPEN	-1.06	**-0.06**	0	0.00	-6.86***
HICPFPR	-0.60	**0.40**	0	0.00	-3.93***
HICPFU	-1.17	**-0.17**	0	0.00	-6.99***
HICPIG	-0.56	**0.44**	0	0.00	-3.78***
HICPSERV	-0.85	**0.15**	0	0.00	-5.70***
HICPSERVXADCT	-0.86	**0.14**	0	0.00	-6.04***

Notes:
Inflation variables are denoted as explained in Table 1.

1) (ρ-1) is the coefficient estimated by the ADF and PP tests based respectively on equations (6) and (7)
2) ρ is the sum of autoregressive coefficients derived indirectly from (ρ-1)
3) The lag length in the ADF test is selected automatically by the Schwarz Info Criterion (BIC), max length=5
4) The default options of EViews are used for the spectral estimation method and the bandwidth in the PP test
5) P-value is the probability associated with the null hypothesis of a unit root in the respective inflation series
6) t-statistics is the statistics of the coefficient (ρ-1) in the OLS equation estimated by the tests
7) MacKinnon's critical values are used for the result on the t-statistics
8) ***, ** and * denote rejection of the non-stationarity null at the 1%, 5% and 10% respectively
9) Numbers in bold indicate that it is possible to reject statistically that ρ=1

[1]ADF and PP unit root tests are applied to the deviation of inflation from its time-varying mean
The time-varying mean of an inflation series is proxied by HP filter-based trend component
The smoothing parameter is set to 1600, the standard number for quarterly data
Inflation series are seasonally-adjusted with Census-X12 procedure, using additive adjustment
ADT (administered prices and tobacco) inflation is not seasonally adjusted
No estimates of inflation persistence are made for ADT

Table 4. Estimates of inflation peristence /constant mean/
Annual inflation rates
Sample period: 1998:q1 - 2007:q4, Adjusted sample: 1999:q1 - 2007:q4

Equation (2) /ADF unit root test, intercept included/

Inflation variable	ρ-1	ρ	Lags	P-value	t-statistics
aggregate level					
CPI/HICP	-0.39	**0.61**	1	0.02	-3.35**
HICPXADT	-0.30	**0.70**	1	0.04	-3.00**
HICPXEUN	-0.34	**0.66**	1	0.01	-3.77***
HICPXEFADT	-0.02	0.98	1	0.91	-0.29
ADT	-0.24	0.76	1	0.11	-2.57
disaggregate level					
HICPEN	-0.47	**0.53**	3	0.01	-3.80***
HICPFPR	-0.29	**0.71**	1	0.04	-3.00**
HICPFU	-0.32	**0.68**	0	0.04	-3.04**
HICPIG	-0.06	0.94	1	0.76	-0.96
HICPSERV	-0.05	0.95	1	0.68	-1.17
HICPSERVXADCT	-0.18	**0.82**	0	0.00	-6.16***

Equation (5) /PP unit root test, intercept included/

Inflation variable	ρ-1	ρ	Lags	P-value	t-statistics
aggregate level					
CPI/HICP	-0.26	0.74	0	0.13	-2.44
HICPXADT	-0.19	0.81	0	0.19	-2.26
HICPXEUN	-0.17	0.83	0	0.26	-2.05
HICPXEFADT	-0.02	0.98	0	0.81	-0.80
ADT	-0.17	0.83	0	0.23	-2.15
disaggregate level					
HICPEN	-0.27	**0.73**	0	0.07	-2.76*
HICPFPR	-0.06	0.94	0	0.67	-1.18
HICPFU	-0.32	**0.68**	0	0.04	-3.03**
HICPIG	0.02	1.02	0	0.84	-0.67
HICPSERV	-0.14	**0.86**	0	0.07	-2.79*
HICPSERVXADCT	-0.18	**0.82**	0	0.00	-5.06***

Notes:

1) (ρ-1) is the coefficient estimated by the ADF and PP tests based respectively on equations (2) and (5)
2) ρ is the sum of autoregressive coefficients derived indirectly from (ρ-1)
3) The lag length in the ADF test is selected automatically by the Schwarz Info Criterion (BIC), max length=3
4) The default options of EViews are used for the spectral estimation method and the bandwidth in the PP test
5) P-value is the probability associated with the null hypothesis of a unit root in the respective inflation series
6) t-statistics is the statistics of the coefficient (ρ-1) in the OLS equation estimated by the tests
7) MacKinnon's critical values are used for the result on the t-statistics
8) ***, ** and * denote rejection of the non-stationarity null at the 1%, 5% and 10% respectively
9) Numbers in bold indicate that it is possible to reject statistically that ρ=1

Table 4A. Estimates of inflation peristence /time-varying mean/[1]
Annual inflation rates
Sample period: 1998:q1 - 2007:q4, Adjusted sample: 1999:q1 - 2007:q4

Equation (6) /ADF unit root test, no intercept included/

Inflation variable	ρ-1	ρ	Lags	P-value	t-statistics
aggregate level					
CPI/HICP	-0.43	**0.57**	1	0.00	-3.60***
HICPXADT	-0.35	**0.65**	1	0.00	-3.28***
HICPXEUN	-0.43	**0.57**	1	0.00	-4.78***
HICPXEFADT	-0.20	**0.80**	1	0.04	-2.04**
ADT	-0.42	**0.58**	1	0.00	-3.70***
disaggregate level					
HICPEN	-0.39	**0.61**	0	0.00	-3.19***
HICPFPR	-0.39	**0.61**	1	0.00	-4.23***
HICPFU	-0.34	**0.66**	0	0.01	-2.88**
HICPIG	-0.14	**0.86**	1	0.06	-1.88*
HICPSERV	-0.18	**0.82**	1	0.10	-1.62*
HICPSERVXADCT	-0.44	**0.56**	0	0.00	-4.80***

Equation (7) /PP unit root test, no intercept included/

Inflation variable	ρ-1	ρ	Lags	P-value	t-statistics
aggregate level					
CPI/HICP	-0.30	**0.70**	0	0.01	-2.70***
HICPXADT	-0.24	**0.76**	0	0.01	-2.60**
HICPXEUN	-0.24	**0.76**	0	0.02	-2.28**
HICPXEFADT	-0.19	**0.81**	0	0.03	-2.12**
ADT	-0.29	**0.71**	0	0.01	-2.81***
disaggregate level					
HICPEN	-0.39	**0.61**	0	0.00	-3.44***
HICPFPR	-0.18	**0.82**	0	0.10	-1.63*
HICPFU	-0.34	**0.66**	0	0.00	-2.93***
HICPIG	-0.09	**0.91**	0	0.08	-1.71*
HICPSERV	-0.25	**0.75**	0	0.01	-2.50**
HICPSERVXADCT	-0.44	**0.56**	0	0.00	-4.52***

Notes:

1) (ρ-1) is the coefficient estimated by the ADF and PP tests based respectively on equations (6) and (7)
2) ρ is the sum of autoregressive coefficients derived indirectly from (ρ-1)
3) The lag length in the ADF test is selected automatically by the Schwarz Info Criterion (BIC), max length=3
4) The default options of EViews are used for the spectral estimation method and the bandwidth in the PP test
5) P-value is the probability associated with the null hypothesis of a unit root in the respective inflation series
6) t-statistics is the statistics of the coefficient (ρ-1) in the OLS equation estimated by the tests
7) MacKinnon's critical values are used for the result on the t-statistics
8) ***, ** and * denote rejection of the non-stationarity null at the 1%, 5% and 10% respectively
9) Numbers in bold indicate that it is possible to reject statistically that ρ=1

[1]ADF and PP unit root tests are applied to the deviation of annual inflation from its time-varying mean
The time-varying mean of annual inflation series is proxied by HP filter-based trend component
The smoothing parameter is set to 1600, the standard number for quarterly data

Table 9

FREQUENCY OF PRICE CHANGES IN BULGARIA IN THE VARIOUS MONTHS, ON AVERAGE PER MONTH (JAN/2000-DEC/2007, %)

Overall CPI/HICP index/representative product prices	Share of unchanged prices in all prices	Share of price changes in all all prices	Share of price increases in all all prices	Share of price decreases in
January	21	79	47	32
February	34	66	40	27
March	36	64	36	28
April	36	64	34	30
May	37	63	33	31
June	38	62	29	33
July	36	64	34	29
August	37	63	37	26
September	33	67	48	19
October	35	65	47	19
November	34	66	46	20
December	37	63	46	18
max	***38***	***79***	***48***	***33***
average	*34*	*66*	*40*	*26*

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