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Determinants of Inflation in a Transition Economy: The Case of Ukraine

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Abstract

The views expressed in this Working Paper are those of the author(s) and do not necessarily represent those of the IMF or IMF policy. Working Papers describe research in progress by the author(s) and are published to elicit comments and to further debate.

This paper examines determinants of inflation in Ukraine during 1993–2002 in a cointegrating framework. Two basic theoretical models—a markup and a money market model—are tested. While broad money is cointegrated with the CPI for the whole sample and for early subsamples, the cointegration ceases to be statistically significant between 1996–2002, in part because of strong remonetization. The mark-up model offers a more consistent and well-fitting overall framework for 1996–2002 data, pointing inter alia to a greater role of administered prices in the CPI within a fairly mainstream inflation process. The "long-term" monetary transmission mechanism operates through the exchange rate and wages, but broad money clearly enters short-term inflation determinants. Prudent macroeconomic policies, grain harvests, and administrative decisions explain the sharp decline of inflation over 2000–2002.

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I. INTRODUCTION

Macroeconomic stabilization has been an important task of economic policy in transition economies. Although this objective was eventually achieved in most countries of Central and Eastern Europe, there were a number of reversals, notably in the context of the 1998 Russian financial crisis. Furthermore, the swings in money demand—mostly in the form of remonetization of late—have been substantially larger than expected, often defying exante inflation projections. Thus, questions continue to persist regarding: (i) the underlying determinants of inflation and their evolution over time; (ii) the durability of stabilizations; and (iii) the degree of tightness of macroeconomic policies that would be appropriate in various circumstances. This paper will concentrate on the first question, recognizing that such an analysis would throw light on the other two.

The determinants of inflation in transition economies have been extensively studied in the academic literature, initially in cross-section or panel regressions across a number of transition economies. On the basis of 1992–95 panel data, Fischer, Sahay, and Vegh (1998) find that fixed exchange rates, lower fiscal deficits, and a number of structural reform variables are associated with lower inflation rates. Corey, Mecagni, and Offerdal (1998), in an analysis of the effects of relative price changes on inflation, show that, while money and wage growth were the most important determinants of inflation, relative price variability had a sizable impact during the high inflation associated with price liberalization, with a more modest effect thereafter. The relative price variability was also emphasized in DeMasi and Koen (1997).

With respect to individual transition economies, an analysis of inflation determinants largely concentrated on Russia. Earlier studies (Koen and Marrese (1995), Hoggarth (1996), Korhonen (1998)) were based on fairly short time series of data and regressed inflation (or prices) on past growth of broad money, and sometimes on other variables. The common conclusion was that Russian inflation clearly had monetary roots, with the time lag between growth of a monetary aggregate and inflation increasing from about 3 months in the early years of transition to 6 months subsequently. A more recent study by Nikolic (2000) based on a somewhat longer time series, allowed more sophisticated tests, as well as breaking the sample into subsamples. That study supported the main conclusion of the earlier papers, although a substantial weakening of the money-inflation relationship was found, with the break occurring about 30 months into the transition.

This paper offers an additional contribution to the analysis of inflation determinants in a single transition economy. Its focus is Ukraine, a country with similar structural characteristics to Russia, but with differences with respect to initial conditions, inflation history, and conduct of economic policies. Inflation rates in Ukraine in the early phase of the reform process were much higher than in Russia and most other transition countries. In fact, in the eyes of many observers, high inflation was a defining feature of Ukraine's experience (see Rugerone (1996)). However, as distinguished from some other transition economies, Ukraine's inflation subsided rapidly following a serious attempt at mainstream stabilization. Otherwise Ukraine's inflation experience has been similar to that of other transition economies, and could illustrate many of the general phenomena observed there.

There have been relatively few published empirical studies on Ukraine's inflation. Banaian et al. (1998) found a strong link between money and prices during the high-inflation period of 1993–96. Two recent papers, Piontkivsky (2002) and Zholud (2002), have investigated inflation between mid-1990s and 2001 in an error-correction framework and by OLS for differenced time series, respectively. Both studies conclude that the role of money in the inflation process became insignificant, although their methods have a number of limitations. Piontkivsky assumes, rather than tests for, cointegration, while Zholud's framework does not account for information embedded in the levels of variables. This paper goes further and conducts unrestricted tests of cointegrating relationships between CPI and other variables in a VAR framework due to Johansen (1988), allowing more powerful tests and systematic treatment of monetary and structural factors. In addition, it compares the cointegrating properties of several alternative frameworks available in the literature on inflation.

The remainder of the paper is structured as follows. Section II provides a descriptive history of inflation in Ukraine. Section III outlines general methodology and data issues. Sections IV and V present the main long-term and short-term results, respectively. Section VI discusses several policy issues related to the recent inflation debate during 2002. Section VII offers some concluding observations.

II. AN OVERVIEW OF UKRAINE'S INFLATION EXPERIENCE

Inflation developments in Ukraine following its independence in late 1991 can be roughly divided into four periods (Figure 1). First, there was a period of (gradual) price liberalization and very high inflation, at times crossing the classical threshold of hyperinflation, from 1992 to late 1994. Then there was a process of disinflation from late 1994 to August 1998, as annual CPI growth eventually dipped below 10 percent between late 1997 and mid-1998. Third, in the aftermath of the 1998 Russian financial crisis there was a moderate relapse of inflation, which hovered at a rate of 20–30 percent per annum for a couple of years. Finally, inflation subsided rapidly and unexpectedly to just 6 percent in 2001 and further into a small deflation in 2002. The remainder of this section will briefly survey Ukraine's inflation history and the underlying policy issues, as they appear from an economist's perspective.

Figure 1. Inflation in Ukraine: percent per month 1992-2002
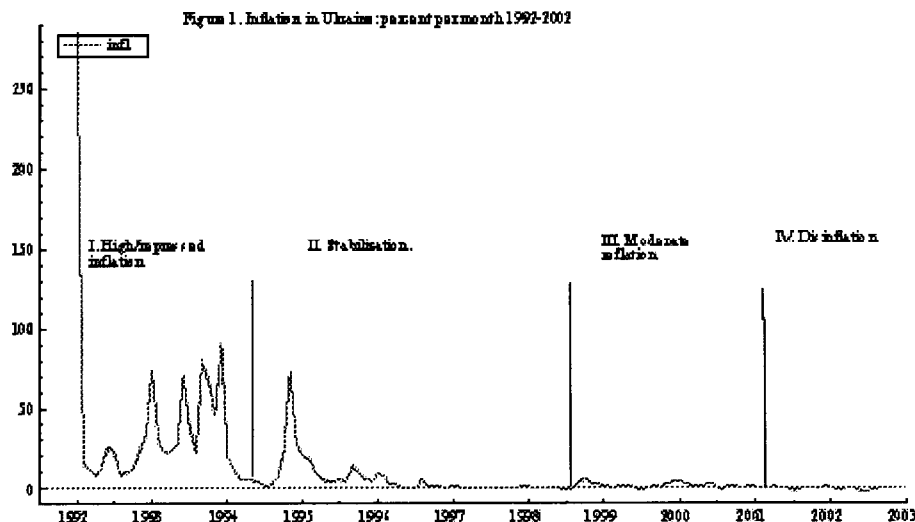
Source: State Statistics Committee of Ukraine.

The initial period of very high inflation was the inevitable outcome of lax macroeconomic policy under a difficult external and domestic environment. The collapse of the former Soviet Union and a consequent move toward the market was accompanied by price liberalization[2] and heavy macroeconomic and policy pressures in the Baltics, Russia and other countries of the former Soviet Union (BRO). In the event, Ukraine pursued fiscal and monetary policies that passively accommodated the imbalances resulting from the prior accumulation of "monetary overhang," slumping revenue, and the rush to maintain the real value of wages, subsidies, and other expenditures. Budget deficits (including noncash and quasi-fiscal activities) averaged some 20 percent of GDP, with almost full monetization of enormous financing gaps. Monetary aggregates, wages, and the exchange rate were increasing (i.e., depreciating) rapidly in nominal terms, but fell sharply in real terms. In 1993 alone, inflation exceeded 10,000 percent, while nominal wages, broad money, and the exchange rate increased by 3,850, 1,900, and 3,350 percent, respectively. As the economy was demonetized, there was little progress in structural reform, since high inflation diverted society's attention and energies from other economic problems.

The first serious stabilization effort was attempted in October 1994 in conjunction with a number of first-generation structural reforms. Most remaining price controls were lifted, export and import quotas abolished, and the exchange rate unified. At the center of stabilization was fiscal adjustment, whereby the overall budget deficit (including estimated noncash operations) was reduced from some 15 percent of GDP in 1994 to 5 percent of GDP in 1995 (see Orsmond (1997)). As a result, inflation fell to 400 percent in 1994 and about 180 percent in 1995, notwithstanding the substantial upward adjustments in administered prices. The nominal exchange rate stabilized in late 1995, and did not change significantly for almost three years, supported by an exchange rate band of Hrv 1.8–2.25 per US$1. The introduction of the new currency—the hryvnia—in September 1996, also helped stabilize expectations. Exchange market stability contributed to a further reduction in end-period inflation to 40 percent in 1996 and just 10 percent in 1997.

In early 1998, inflation continued to fall, but—paradoxically—Ukraine found itself in a precarious financial position. First, the fiscal adjustment had suffered a substantial setback, as the budget deficit increased from 3.2 percent of GDP in 1996 to 5.6 percent in 1997. While Ukraine was temporarily able to finance the deficit in a noninflationary way through domestic treasury bills and in international debt markets, fiscal sustainability was increasingly questionable due to the debt's short maturity and high interest rates, prompting a reversal in capital inflows. Second, Ukraine's external current account—after a prolonged real appreciation within the exchange rate band—deteriorated rapidly with the fiscal expansion and the export slowdown caused by difficulties in Russia and Asia. The current and capital account deficits led

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[2] In Ukraine, price decontrol occurred more gradually than in many BRO countries, although this curbed aggregate prices only partially, leading to a strange coexistence of shortages and high inflation (Lane et al. (1994)).

to a substantial reserve loss, which contained monetary expansion—and hence inflation—but could not go on indefinitely.

The situation became clearly untenable with the massive August 1998 default and devaluation in Russia. With foreign reserves falling to just over a week of imports, the authorities let the currency depreciate by more than 50 percent, modifying the limits of the band. Simultaneously, exchange restrictions were introduced by the National Bank of Ukraine (NBU) to prevent a further slide in the rate due to "speculation." Importantly, all this was supported by a tightening of fiscal and monetary policies. The measures succeeded in containing the inevitable inflationary spike. Initially, inflation picked up in late 1998, bringing the end-1998 12-month rate to 20 percent, but in the first quarter of 1999 slowed to 1 percent per month compared with over 4 percent in the previous quarter. As the authorities regained control over Ukraine's financial situation, a floating exchange rate regime was adopted in March 1999 to facilitate external adjustment under a very weak reserve position. At this time, most exchange restrictions were removed, causing the rate to drop further by some 15 percent.

In the ensuing two years, inflation remained in the "moderate" range of 20–30 percent per annum, partly reflecting external and structural imbalances. In 1999, prices had to absorb two further, albeit moderate (15–20 percent per episode), depreciation spurts.[3] With significant cumulative depreciation boosting Ukraine's external competitiveness, the nominal exchange rate stabilized in early 2000. However, inflation did not subside in 2000 in part because of several adjustments in administered prices for communal services and staples. The former were increased toward cost recovery levels in the first half of 2000, while poor harvests of 1999–2000 caused a substantial increase in food, especially bread, prices. Thus, despite many signs of macroeconomic stability—such as the beginning of economic growth, fiscal consolidation, exchange rate stability, orderly debt restructuring, and accumulation of reserves—12-month inflation at end-2000 rose to 26 percent, the highest level since 1996.

As the extent of inflation in 2000 remained a puzzle even after accounting for administrative price adjustments, some observers linked it to substantial nonsterilized foreign exchange purchases by the NBU. Indeed, nominal broad money had expanded by 45 percent during 2000, possibly running into capacity constraints manifested during the economic recovery.[4] This argued for a more cautious monetary policy stance, involving greater sterilization of foreign exchange inflows, and even some nominal exchange rate appreciation, to be balanced against the continued need to build up external reserves. However, there was substantial uncertainty both with respect to the extent of future inflows and the expected money demand.

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[3] The first significant depreciation (mid-1999) followed a shortage in the market for petroleum products. Another followed a general loosening of macroeconomic policies in the last quarter of 1999.

[4] While there was evidence of substantial **physical** excess capacity (idle Soviet-era plants), excess viable ("value-creating") capacity under a market economy appeared much smaller.

In 2001–02, the inflow of foreign exchange continued, while the NBU opted for very limited sterilization of the rapid increase in the money supply. Paradoxically, inflation subsided to just 6 percent at end-2001. This "reverse puzzle" of low inflation in 2001 (which was reinforced in 2002) had some explanations, including: (i) a fall in agricultural prices after the good 2001–02 harvests; (ii) an increase in the share of cash transactions, thereby contributing to the remonetization and to a direct reduction in reported prices as premiums for noncash transactions evaporated; (iii) delays in the adjustment of administered prices for services despite indications that these were further slipping below cost recovery level. Still, there were doubts whether these reasons could fully account for the excellent inflation performance in 2001–02, which coincided with rapid growth in real output, thereby apparently refuting the conjecture that capacity constraints had been "biting" to date.

In sum, unlike its poor growth record throughout the 1990s, Ukraine's inflation performance since the mid-1990s was relatively successful, even compared to the more advanced reformers in Eastern Europe.[5] By and large, inflation responded quickly to the initial stabilization effort (Figure 2), and starting from late 1996 basically hovered in the "moderate" range of 15–30 percent (dipping further toward the end of the sample period). However, this moderate inflation often behaved differently from what policymakers could expect, and it would thus be interesting to explore with more rigor whether these outcomes were puzzling indeed.

Figure 2. Ukraine: log of 12-month inflation 1993-2002
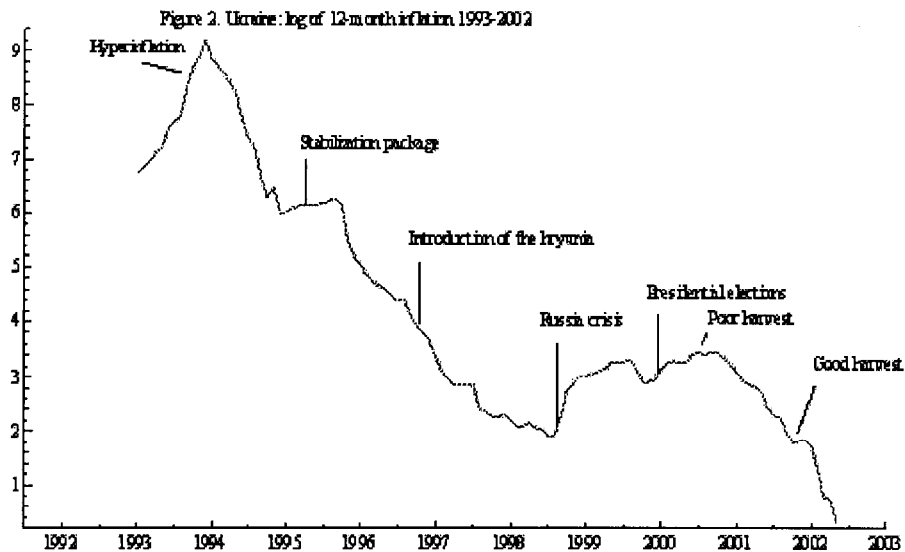
Source: State Statistics Committee of Ukraine.

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[5] For example, Poland's inflation remained at or above the moderate range years after the comprehensive stabilization effort in the early 1990s (see Cottarelli and Doyle (1997)).

III. THE MODELING FRAMEWORK

A. General Approach

It transpires from the previous discussion that Ukraine after independence went through very diverse periods in terms of inflation rates per se, the nature of imbalances that were causing it (domestic, external), exchange rate regimes, and sectoral factors. Thus, to the extent there is a unifying framework for explaining inflation during the whole period or a substantial portion thereof, it would need to test and possibly incorporate the effects of disparate factors and recognize that some of the structural parameters may not be stable over time. The analysis would be further complicated because the timing of some decisions (administered price increases) is often politically motivated.

Theoretically, the inflation process is usually modeled as: (i) an interaction of supply and demand for money (Jonsson (1999)); (ii) a markup pricing mechanism (De Brouwer and Ericsson (1998)); or (iii) particular factors affecting external (equilibrium exchange rates) or domestic (expectations) disequilibria pressures. The more sophisticated applied approaches involve various combinations of the above, with some allowance for the structure of the real economy, often emphasizing aspects of energy pricing or an explicit allowance for administered prices.

This paper intends to capture the key sources of inflationary pressure in Ukraine, in part relying on the previous section's exposition of the most important stylized facts. In particular, the most relevant stylized facts to address appear to be the following:

- Impact of domestic financial disequilibria that were apparent in the early transition (monetization of budget deficits and other reasons for "excessive" growth of the money supply).
- Role of external disequilibria (currency crises, external shocks).
- Pronounced seasonality of Ukraine's economy, including that of prices and production.
- Convergence of Ukraine's price level toward that of the world economy.
- Allowance for the timing of administered price increases.
- Allowance for the output gap and other factors that may be behind a significant remonetization of the economy.

It is also clear that the main potential explanations differ in terms of the nature of their impact on inflation, encompassing temporary and more prolonged phenomena. For example, some factors are clearly short-term (episodes of shocks and financial disequilibria), some are "medium-term," such as convergence of domestic prices to world price levels that characterizes most transition economies, while some would be longer-term, such as the observance of PPP relationships for the subset of traded goods after current account transactions are sufficiently liberalized. In these circumstances, an eclectic cointegration framework would be called for to separate key short-term and longer-term determinants of inflation. This approach needs to be

fairly general in order not to prejudge the causal links in an economy characterized by substantial macroeconomic disturbances and structural change.

B. Data

The data are monthly, although this implies a cost as the quality of most quarterly or annual series—particularly those for output—is usually better than the monthly series. However, since only about 10 years have passed after Ukraine gained independence, annual or quarterly data would not provide a sufficient number of observations for a meaningful statistical analysis. Furthermore, until the mid-1990s, Ukraine's economic data across the board were characterized by huge variability and poor quality. Since then, there has been substantial, albeit incomplete, progress with data quality. All primary information used in this study is publicly available.

Price trends in Ukraine can be assessed on several measures. The main indicator is the Laspeyers-based consumer price index (CPI) calculated by the State Statistics Committee (SSC) on a monthly basis. The index comprises a basket of goods and services tracked in various locations across Ukraine. The item sample and weights for the index have been updated annually based on the results of household budget surveys. Over time, the index has been simplified with foreign technical assistance and without much loss of geographical representation. The number of product items in the sample has been reduced from 425 to about 270, while the number of urban areas (rayons) from which the data are collected was also reduced from about 700 in 1995 to about 400 lately. These changes have allowed a substantial reduction in the administrative burden involved in the calculation of the CPI.

The index is decomposed into three subcomponents: (i) food, (ii) nonfood items and (iii) services. For each of these, a separate index is available (Figure 3). The weights of the subcomponents in the aggregate CPI have varied somewhat. For example, they ranged from 50–70 percent for food, which has been the main component throughout. The shares of each of the other two components have ranged between 15 and 30 percent, depending on the time period. This time-variability may create some difficulties with statistical inferences from the CPI data. An additional complication is the large weight of the volatile—and strongly seasonal—food prices, which depend on a number of stochastic supply-side factors. As can be seen from Figure 3, services prices have had a number of discrete adjustments without a strong seasonal pattern, but rather based on some adhoc considerations. It can also be seen from the lower panel of Figure 3 that services prices grew more quickly than other components of the CPI throughout the post-independence period. Prices for non food items have generally followed exchange rate changes, as the weight of tradables in this component is fairly large.

The CPI coverage has a couple of important specificities. First, it is an urban CPI and does not cover a third of the population living in rural areas. Second, the consumption basket is compiled on the basis of the household budget survey, which usually does not capture the underreported

Figure 3. Components of the CPI rates of change (1996-2002) and levels (1992-2002)
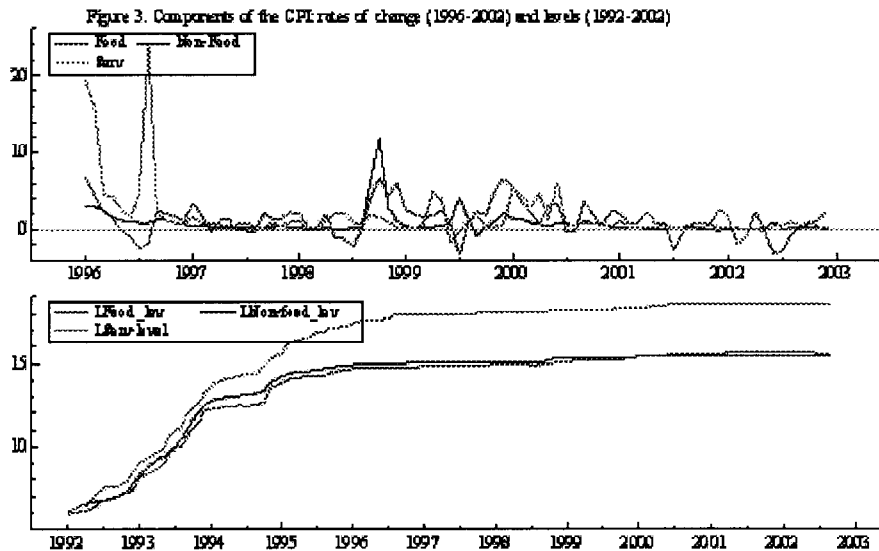
Source: State Statistics Committee of Ukraine.

high end of the market.[6] Given the substantial income differentiation in Ukraine, this may have seriously affected the "economic" coverage of the CPI. For example, the annual size of the car market in Ukraine is plausibly estimated by the auto dealers at around Hrv 13 billion, which accounted for over 10 percent of total consumer expenditure in Ukraine in 2001. However, the share of automobiles in the CPI basket has been smaller by a factor of 100. Lately, there has been a general feeling that inflation declined by a lot less than indicated by the CPI during 2002, prompting doubts as to whether the latter truly reflected the situation with price growth.[7]

Another important price measure in Ukraine is the producer price index (PPI). These data are also monthly and based on sectoral information compiled from the industry surveys. The PPI is also rebased annually using the information from the surveys. However, the index has had a number of methodological problems throughout the whole transition process. For example, there have been significant delays in updating the weights for many years, which has affected the quality of the index. There have also been problems in chaining the index and in recording correct average sales prices at the time of production, including both exported output and output sold to resident units. Despite certain methodological problems, the PPI has been an important gauge of inflation processes in Ukraine, especially because of their CPI coverage issues.

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[6] Richer people often refuse to participate in these voluntary surveys as they are not sure of their confidentiality. Also, the fixed remuneration for survey participants is a bigger incentive for poor people to participate in these surveys.

[7] For example, some Ukrainian consumer associations disputed official numbers, arguing that the effective increases in prices were much larger than declared.

Other macroeconomic variables compiled in Ukraine on a monthly basis include: (i) output data (nominal and real) and in particular GDP, industrial and agricultural production indexes (these output indicators have had significant methodological problems throughout the transition period, and most good measures have been available only after the mid-1990s); (ii) wage data (although these are only accrual-based and exclude the impact of wage arrears); (iii) exchange rate data (these are relatively reliable, except during the period of heavy exchange restrictions from September 1998 to March 1999); the U.S. dollar exchange rate has in practice been dominating all other exchange rates, although those versus the Russian ruble and the euro also play some role; (iv) monetary data for narrow and broad money aggregates; (v) fiscal deficit data (however, the quality of the data, especially monthly, is poor for much of the sample, partly because of extensive quasi-fiscal activities); (vi) interest rate data (although with missing observations for many series); (vii) data on noncash transactions, such as interenterprise arrears and barter transactions in industry (the data on barter only became available in the second quarter of 1997).

C. Derivation of the Price Equation(s)

The overall price level (P) is a weighted average—expressed in log-linear form (denoted by lower-case letters)—of the price of tradable and nontradable goods, where α represents the share of non-tradable goods in total expenditure:

$$p = \alpha(p_N) + (1-\alpha)p_T \tag{1}$$

The price of tradable goods is exogenously determined in the world market and, in domestic currency terms, can be expressed as a function of foreign prices and the exchange rate:

$$p_T = e + p_f \tag{2}$$

Both an increase (i.e., depreciation) in the exchange rate and foreign prices will lead to an increase in the overall price level. Furthermore, given that world inflation was relatively minor in the observed period compared to domestic inflation, the last term in this equation could be normalized and therefore dropped, without significant loss of explanatory power, yielding:

$$p_T = e \tag{2a}$$

The price of nontradable goods is assumed to be set in two alternative ways. On the one hand, it could be an outcome of the money market equation, in which demand for nontradable goods is related to the economy-wide demand:

$$p_N = \beta(m^s - m^d) + \varphi p_{adm} \tag{3}$$

where m^s represents the nominal stock of money, m^d is demand for real balances, and P_{adm} is the term that accounts for the administrative component of domestic prices, since the extent and timing of some increases are determined by the decisions of the government and not by market forces.

The demand for real balances is assumed to be a function of all or some of the following variables:

$$m - p = f(\overset{+}{y}, \overset{-}{Dp^e}, \overset{-}{i}, \overset{-}{b}) \tag{4}$$

where y is real income, Dp is the (expected) rate of inflation, i is the level of an interest rate, b is the share of barter (or, if available, noncash) transactions in the economy. To the extent an explicit money demand equation could not be specified, other approaches to price determination are available.

Alternatively, the price of nontradable goods can be represented in terms of a markup model:

$$p_N = \mu + \kappa w + \gamma p_{adm} \tag{5}$$

where w is the domestic wage level, which approximates the cost of labor, and μ is some markup coefficient (see De Brouwer and Ericsson (1998)). Note that the coefficient that adjusts the nontradables price level to changes in administrative prices may well be different in (3) and (5), given the obvious differences in the conceptual model of price determination. The rationale for equation (5) is reinforced by the fact that the consumption basket in Ukraine mostly relates to expenditure by wage earners.

Thus, the following testable equations can be considered for the aggregate price level:

(i) The **markup model** is obtained by combining (1), (2a) and (5):

$$p = \xi + vw + \rho p_{adm} + \delta e \tag{6}$$

where $\xi = \alpha\mu$ is a positive constant and $v = \alpha\kappa$; $\rho = \alpha\gamma$; and $\delta = 1-\alpha$ are all positive fractions. Linear homogeneity would imply the sum of three fractions would be equal to unity.

(ii) The **money market** model would imply the following broad relationship:

$$p = f(\overset{+}{m}, \overset{-+}{y}, \overset{+}{Dp}, \overset{+}{i}, \overset{+}{b}, \overset{+}{e}, \overset{+}{p_{adm}}, ...) \tag{7}$$

For a fuller examination of (7), separate estimates of a robust money demand would also be desirable. Note that the relationship between prices and output can be ambiguous. It could be negative because of the money demand considerations and positive because of the "output gap"

reasons. However, it is very difficult to estimate the output gap in a transition economy directly, so this ambiguity needs to be kept in mind during the empirical tests.

IV. LONG-TERM RELATIONSHIPS

A. Integration

The available data were first tested for the order of integration. The results are presented in Tables 1 and 2. All measures of inflation appear stationary processes according to the Augmented Dickey-Fuller (ADF) tests, at the 1 percent significance level (while the measures of the price level are integrated of the order 1). The basic integration result is insensitive to the period chosen, although inclusion of the period of high inflation of the early 1990s into the sample weakens the significance of the tests somewhat.

Table 1. Integration Properties of Inflation Measures, 1996–2002

	Dp (CPI)	Dp (food)	Dp (non-food)	Dp (services)	Dp (producer)
ADF	-6.18**	-5.13**	-5.88**	-7.28**	-4.63**
Specification[8]	C (1)	C (1)	C (1)	C (1)	C (1)
DW	1.94	1.94	1.95	1.99	2.01

** and * denotes significance at the 1 percent and 5 percent levels, respectively.

Table 2. Tests for the Order of Integration of Other Selected Variables (1996–2002).

		y	W	E	Brd.M	Base.M	i-lend	i-dep	i-int
Levels	Adf	-2.36	-1.00	-1.45	1.75	0.39	-4.40**	-4.19**	-4.63**
	Dw	1.95	2.31	1.78	2.18	2.01	1.98	1.92	2.07
Changes	Adf	-6.30**	-5.78**	-11.78**	-8.00**	-7.77**	-5.81**	-7.28**	-5.46**
	Dw	2.02	1.94	2.03	1.98	2.04	2.01	2.01	1.88

Table 2 shows that such variables as real output, wages, the exchange rate and monetary aggregates are stationary in their first differences, while lending, deposit and interbank rates are already stationary in levels.

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[8] Indicates whether a constant term (C) has been introduced and the lag length of the dependent variable (in parentheses). The robustness of this result was checked by introducing more lags, and the basic integration results did not change.

B. Cointegration

The markup model

A long-run cointegrating relation between consumer prices, wages, the exchange rate (versus the U.S. dollar) and services prices (p_{serv}), as a proxy for administrative prices, was estimated based on the Johansen (1988) VAR procedure, using six lags, a constant and seasonal dummies.[9] The sample period was initially chosen starting from January 1993, as the year 1992 is characterized by particularly poor data, a problem exacerbated by pervasive price controls at the time. The main results are presented in Tables 3 and 4.

Table 3. Cointegration Analysis of Equation (6), 1993(1)–2002(9): Eigenvalues

Ho: rank = p	λ	λmax	95%	λtrace	95%
p == 0	0.48	77.5**	27.1	111.9**	47.2
p <= 1	0.17	21.6*	21.0	34.4*	29.7
p <= 2	0.10	11.7	14.1	12.8	15.4
p <= 3	0.01	1.1	3.8	1.1	3.8

Table 3 implies the existence of two statistically significant long-term cointegrating vectors, one at 1 percent significance level and the other at 5 percent. As is clear from Table 4, both these vectors (top two rows) roughly conform to the theoretical relationship in equation (6): all variables in these equations have the right signs and the sum of normalized coefficients for the exchange rate, wages, and administrative prices is equal to 0.7 in the first equation and to 0.91 in the second equation. Restricted cointegration analysis does not reject the hypothesis of unity for either equation. The individual coefficients are pretty similar in both equations, with the services prices playing the largest role, although in the first equation their (long-term) impact on the overall price level is smaller. The exchange rate has the smallest long-term coefficient of 0.11–0.13, although it roughly corresponds to the share of tradables in Ukraine's consumer basket, which has been estimated between 10 and 15 percent in the period under consideration.

The same cointegrating relationship was tested with the Engle-Granger procedure, and the following static equation was obtained, which confirms qualitative results in Table 4, although with somewhat different coefficients.

$$p = 7.4 + 0.45w + 0.33p_{serv} + 0.17e \tag{8}$$

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[9] The Johansen VAR procedure is generally considered better than alternative methods to assess cointegration, partly because it is less dependent on theoretical "priors," allows simultaneous evaluation of multiple cointegration vectors, and is more powerful in small samples than, say, the Engle-Granger procedure. The latter will sometimes be used in the course of the paper to check the robustness of some results.

Table 4. Cointegration Analysis of Equation (6), 1993(1)–2002(9): Normalized Matrix of Linearly Independent Vectors

Variable				
p**	1.000	-0.109	-0.249	-0.341
e*	-7.489	1.000	1.289	4.551
w	0.906	-0.973	1.000	-0.792
p_{serv}	20.283	-6.823	-17.149	1.000

Since the two significant cointegrating vectors are broadly similar, it is possible that they correspond to the same theoretical relationship, but reflect structural changes in the coefficients over time, perhaps because of the structural break between the high and low inflation environment. Thus, the cointegration properties of the subsamples have been investigated further. Much of the rest of the paper will concentrate on the 1996–2002 period, which has been characterized by relatively stable inflation and better data quality.

Table 5 presents the results of the Johansen VAR procedure for the 1996–2002 data. To get a parsimonious VAR, the number of lags was sequentially reduced from 6 to 2, and this reduction was statistically acceptable.

Table 5. Cointegration Analysis of Equation (6): 1996(2)–2002(9), Eigenvalues

Ho: rank = p	λ	λmax	95%	λtrace	95%
p == 0	0.32	31.0*	27.1	67.2**	47.2
p <= 1	0.25	22.9*	21.0	36.1**	29.7
p <= 2	0.15	12.7	14.1	13.3	15.4
p <= 3	0.01	0.6	3.8	0.6	3.8

Table 6. Cointegration Analysis of Equation (6), 1996(2)–2002(9): Normalized Matrix of Vectors and Feedback Coefficients

Variable	Vectors				Feedback coefficients			
p*	1.00	-0.08	-0.32	-0.45	0.022	-0.002	-0.000	-0.002
e*	-25.32	1.00	0.93	26.97	-0.019	-0.000	-0.002	-0.038
w	84.91	-30.26	1.00	-37.34	-0.229	0.000	-0.000	-0.000
p_{serv}	-3.28	0.84	1.02	1.00	-0.018	-0.008	-0.000	-0.000

It can be seen that there are two significant eigenvectors corresponding to the period of relatively low inflation in 1996–2002 (Figure 4). The first row vector in Table 6 closely corresponds to the relationship posited in equation (6). The sum of the coefficients is quite close to unity, pointing to linear homogeneity. Note that the coefficient for the exchange rate has become even smaller, while that for the prices of services is significantly larger than for the whole sample. The former may indicate that the long-term role of the exchange rate in influencing prices has declined over time, perhaps partly reflecting a significant import

Figure 4. The Significant Cointegrating Vectors of Tables 5-6.
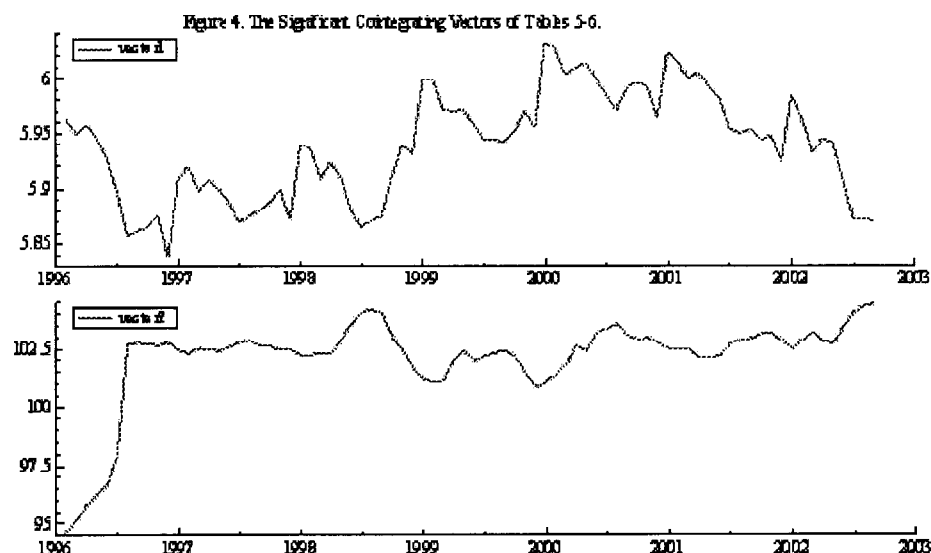
substitution in Ukraine during the 1990s, characterized by the restructuring of the food industry. The high coefficient for prices for services may indicate that, with basic macroeconomic stabilization, the main "long-term" determinant of prices has shifted from "mainstream inflationary pressures" toward increases in the administered (mainly, services) prices from the extremely low levels of the early transition.

The second vector has a similar economic interpretation, although the coefficient on the services prices becomes very high, while those for wages and the exchange rate become extremely small. Two other possible cointegrating vectors are insignificant, but also broadly fitting the general markup assumptions. Interestingly, the values there indicate a much higher long-term coefficient on the exchange rate (0.3–0.4), compared to the significant cointegration vectors.

Since these results may be sensitive to the choice of the sample period, the Johansen VAR cointegration test was run on all monthly samples starting between January 1995 to October 1996, and ending September 2002. The results overwhelmingly indicate presence of cointegration along the lines of the markup equation: in more than 80 percent of cases an equation broadly consistent with the markup equation exhibited cointegration with at least 95 percent significance. Furthermore, in all other cases an equation corresponding to the markup equation was one of the possible cointegrating vectors, and the significance level often approached 90 percent.

To the extent cointegration is established, there is the question on how many cointegrating vectors are statistically significant, which would be crucial for further empirical investigation. Tables 3–6 indicate two such vectors. However, the results are mixed if all monthly samples are examined for the 1995/6–2002 period: sometimes yielding one cointegrating vector, sometimes two, and in some cases even more cointegrating vectors. What is encouraging, though, is that all these vectors are broadly similar, possibly indicating that there is one underlying cointegrating vector, which is somewhat elusive to establish because of substantial data weaknesses. Thus, the analysis in Section IV below is based on a single cointegrating vector in the markup equation.

The money market

One approach for testing the significance of money as a direct determinant of inflation is to simply evaluate a (long-term) relationship between money and prices (Nikolic (2000)). To assess cointegration, the Johansen VAR procedure was applied to the CPI and the broad money aggregate incorporating foreign exchange deposits.[10] Various lag structures were tested. The main results of this exercise for different sample periods are presented in Table 7, using six lags and seasonal dummies.

Table 7. Cointegration of Money and Prices: Significance of the Main Cointegrating Vector

Sample	λ	λmax	95%	λtrace	95%
1993–2002	0.51	79.7**	14.1	80.4**	15.4
1994–2002	0.18	20.8**	14.1	21.2**	15.4
1995–2002	0.12	12.4	14.1	13.4	15.4
1996–2002	0.10	8.3	14.1	8.4	15.4

As appears from the table, the "long-term" money—CPI link is progressively weakened as the sample captures less of the relatively high inflation period of 1993–96. Incidentally, the cointegration relation tested just for the period between 1993–96 is significant at the 1 percent level, indicating cointegration even within this relatively short sample. The absence of a significant long-term money-CPI relationship between 1995/6 and 2002 indicates that, with the attainment of stabilization, the paths of money and prices diverge as the economy is remonetized. The results for the full sample are still significant, partly because they include both "demonetization" and remonetization, and partly because the high variation in the data during 1993–96 apparently dominates relationships in other periods. It is also possible that the money-CPI link was still present during 1995/6–2002, but has to be controlled for the presence of other variables, particularly those that play a role in the remonetization process.

To check the latter conjecture, a "general-to-specific" approach was used to test the role of money as a "long-term" determinant of inflation for the period 1995/6–2002. Thus, the logs of broad money and CPI inflation were tried in a combination with the other I(1) variables in the data set, with as many variables and with as many lags as possible, and with a successive simplification of the VAR models. In general, the results of these exercises have not supported the hypothesis of a significant direct long-term link of the monetary aggregates with CPI inflation between 1996 and 2002. In many tests, the inclusion of a monetary aggregate in the markup model of the previous section substantially weakens the cointegration of the "markup" vector, often rendering it statistically insignificant. Further, the long-term coefficient for the

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[10] It has been revealed in most transition studies (see Nikolic (2000)), that broader monetary aggregates generally bear a closer relation to inflation than narrower aggregates. Some tests with narrower aggregates were also performed but are not reported here.

services prices, and sometimes for broad money itself, awkwardly becomes negative. In other variable combinations, a robust cointegrating relationship involving both money and the CPI could not be quickly detected either.

One possible exception of note is the VAR between money, real output, and the CPI. For most monthly samples starting during 1996 and ending in September 2002, the VAR indicated fairly high eigenvalues consistent with the following possible cointegrating vector:

$$p = 0.2m + 1.5y \tag{9}$$

However, this eigenvalue—and hence cointegration—was statistically insignificant (at least at the 5 percent level) in most samples covering the 1995/6–2002 period. In addition, the zero coefficient on money often cannot be rejected in a restricted cointegration analysis. On the other hand, an ADL estimate of the same model yields very similar results, thereby indirectly supporting equation (9). If so, this points to a positive, albeit weak, long-term influence of money on prices and to a significant long-term impact of output on the CPI, indicating a possible role of the output gap in price determination.

A more structured approach to assessing the money market is to estimate a long-term money demand relation along the lines of equation (4). Besides the output gap, there have been several other indicators that may explain the remonetization of the economy, such as: (i) the ongoing shift from noncash to cash-based transactions; (ii) improving corporate governance both on the part of enterprises and commercial banks; and (iii) a gradual perception of durability of stabilization by the economic agents in the economy. These factors have clearly been present at the later stages of the transition process and appear to have played some role in inducing greater holdings of domestic currency assets.[11]

Several "remonetization conjectures" were thus tested. The Johansen VAR procedure yields the following results for the available I(1) variables.[12] Tables 8 and 9 show that a robust money demand equation appears somewhat elusive for 1996–2002. While there is one statistically significant cointegrating vector, the signs of long-term coefficients do not correspond to economic theory, indicating a negative relationship between real balances and real output if barter is included as another explanatory variable. Only the last cointegrating vector at the

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[11] A similar money demand function was estimated for Russia (Banerji (2002)) that included, inter alia, all of the proposed terms. While many of the estimated coefficients seemed plausible, a *well-specified* model of money demand could not yet be constructed based on the available sample of observations for Russia.

[12] Since inflation and interest rate variables clearly appear to be I(0), we have not included these in the cointegrating VAR, although in other studies (Banerji (2002), Ahumada (1992)) these variables were assumed I(1) in view of mixed evidence. The inclusion of these variables in Table 8 does not appear to strengthen cointegration results.

bottom of Table 8 appears to be broadly consistent with the theoretical money demand relation, but the eigenvalue is very small and statistically insignificant. If barter is excluded from the cointegrating relation, the sign on real output becomes "right," but the potential cointegrating vectors are not statistically significant (see Table 9).

Table 8. VAR Cointegration Analysis of Money Demand, Real GDP, and Barter, 1997(7)–2002(9): Normalized Matrix of Vectors, Two Lags, and Seasonal Dummies

Variable				eigenvalue	λ max	95%
$m\text{-}p^*$	1.000	1.950	0.600	0.295	22.0	21.0
y	0.002	1.000	0.018	0.180	12.5	14.1
b	9.000	-9.705	1.000	0.003	0.2	3.8

Table 9. VAR Cointegration Analysis of Money Demand and Real GDP: Normalized Matrix of Vectors 1996(6)–2002(9)

Variable			eigenvalue	λ max	95%
$m\text{-}p$	1.00	-3.44	0.06	4.7	14.1
y	-1.26	1.00	0.03	2.2	3.8

Imposing further structure through estimating a long-term money demand equation by OLS does not—even remotely—confirm the results generated by the VAR, indirectly pointing to misspecification problems. The coefficients thus obtained (the first two columns of Table 10) do not correspond to the relevant coefficients in Tables 8 and 9, have large standard errors, and often have "wrong" signs with respect to real output. Including other theoretically plausible variables (right columns of Table 10) generally does not yield many interpretable results either, beyond the fact that the "correct" signs are fairly robust for barter and inflation variables.

Table 10. "Long-Run" Estimates of Real Money Demand by OLS, Including Seasonal Dummies, Standard Errors in Parentheses

$m\text{-}p$	1996–2002	1997–2002	1997–2002	1997–2002	1997–2002	1997–2002
y	5.2(2.6)*	0.2(3.7)	-1.3(2.6)	-1.6(2.0)	--	-1.3(1.1)
b	--	-0.66(0.65)	-0.59(0.40)	-0.67(0.42)	-0.44(0.13)	-0.57(0.16)
$i\ dep$	--	--	-0.03(0.05)	--	0.02(0.02)	0.00(0.01)
Dp	--	--	--	-18.6(24.4)	-23.0(17.4)	-8.1(5.5)
Lags	6→1	6→1	4→1	4→1	4→1	3→1
DW	1.89	2.11	2.27	2.16	1.98	2.14

The results in this section suggest that the broad money aggregate has not appeared to be a significant direct "longer-term" determinant of the consumer price level in Ukraine during 1996–2002. This is not altogether surprising in the light of the inflation literature for both transition and other economies. Kuijs (2002) has reached similar conclusions for the Slovakian CPI, while De Brower and Ericsson (1998) documented a much stronger fit of an alternative

markup model for the Australian CPI. Furthermore, the weak observed role of money is consistent with the recent studies on Ukraine. However, several basic caveats to this conclusion are in order.

First, it may well be possible that there was a stronger long-term link between (broad) monetary aggregates and inflation during 1996–2002 than the data suggest, but that the verification of this link is hampered by poor availability or quality of data. In particular, the relevant "monetary" aggregate could well be a measure that includes foreign exchange cash component of the money supply, the data on which are not available. Anecdotal evidence suggests that the share of foreign exchange cash (compared to components of the measured broad money aggregates) may have been decreasing for a significant portion of the 1996–2002 period.[13] If so, though it cannot be detected directly, the long-term monetary transmission mechanism could operate through other channels, say, the exchange rate. Some evidence in favor of this conjecture will be examined below in the context of the markup model.

Second, even if the small role of monetary aggregates is true for the period in question, most economists think that the process of strong remonetization in Ukraine over 1996–2002 is only "long term" in terms of the sample to date, but in all likelihood is "medium-term" in a practical economic sense. Given the current difficulties in coming up with a satisfactory framework underlying this remonetization, its future path is highly uncertain. At that point the strong money-CPI link of type that was documented for 1993–96 could well be reestablished in the (possibly immediate) future.

Finally, the absence of cointegration results by itself does not prove the absence of *any* link between money and prices. Rather, it indicates that a consistent framework for the assessment of the long-term link is not permitted by the data for the particular period of time. While capturing the underlying long-run link may not be possible, a reasonably consistent alternative cointegrating framework may allow to test for money as a short-term determinant of inflation, as the discussion in the next section will show.

V. Short-Run Properties of the Markup Model

A. Weak Exogeneity

Given the evidence of cointegration between wages, consumer prices, the exchange rate, and services prices for 1996–2002, we may proceed toward a more detailed investigation of the above markup model. Most feedback coefficients of Table 6 are pretty small for all cointegrating vectors, which would suggest "weak exogeneity" (see DeBrower and Ericsson (1998)). However, in the first cointegrating vector, which precisely reflects a markup equation, the weak exogeneity assumption does not appear to hold for wages, as the high feedback

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[13] Similarly, the incomplete coverage of the price indices may also obscure the money-CPI link somewhat.

coefficient indicates a strong adjustment of this variable in response to its deviations from the cointegrating vector.

A closer examination shows that the feedback coefficient may be very high for reasons other than violation of the weak exogeneity assumption. VAR-based OLS estimates indicate that, in the equation for wages, monthly dummies are extremely powerful, and dominate any other explanatory variable. This can be explained by the traditional pattern of the payment of wages in Ukraine, characterized by the so-called "13-th" wage bonus paid in December to many employees, after which a sharp downward adjustment in nominal terms follows in January. As a result of this pattern, wages appear to forcibly fluctuate around the cointegrating vector, without necessarily indicating the strong short-term reversion due to endogeneity. This pattern can also be seen in Figure 5.

Figure 5. Wages and the Cointegration Vector.
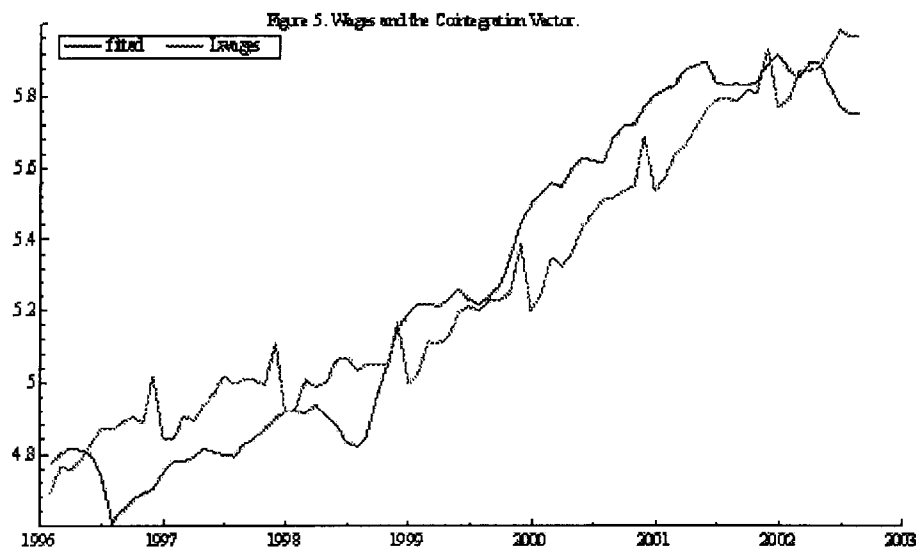
A restricted cointegration analysis of the feedback coefficients was run on the top cointegrating vector in Table 6, in order to see whether a sequence of quantitative assumptions could provide some evidence in favor of weak exogeneity for variables other than the consumer price index. This exercise was unsuccessful, however, as the setting of the feedback coefficient on wages, and sometimes on the services prices, to zero almost always rejected the resulting model. Interestingly, the restricted cointegration analysis that assumes weak exogeneity yields a strongly converging equation that is close to the theoretical markup model and is broadly consistent with estimates of the single-equation model below, but this equation is rejected statistically.[14]

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[14] In particular, the restricted coefficients would be 0.31 for the exchange rate, 0.48 for the services prices, and 0.04 for wages. The fairly low value of the latter may indicate difficulties in properly accounting for the particular seasonality pattern of wages within the VAR.

Nevertheless, weak exogeneity seems reasonable to assume in this particular model. First, the model has a clear theoretical and logical rationale and strong evidence in favor of it was found in other countries (see De Brower and Ericsson (1998)). Second, there is some evidence that the weak exogeneity assumption is rejected "artificially" in a VAR model, due to the particular seasonal pattern of the underlying data: this notably concerns wages, but also to some extent services prices, since many of the adjustments are undertaken at the beginning of every year. Third, the very nature of wages, services prices, and the exchange rate suggests that they can be taken as exogenous for very short periods of time: wages because of the seasonality; services prices because their increases are clustered for seasonal and political reasons, and the exchange rate because for very long spells of time it was either explicitly or implicitly targeted.

In the absence of the weak exogeneity assumption, a proper approach would be to proceed to the short-term dynamics within the VAR framework, with separate equations for variables for which weak exogeneity cannot be confirmed. However, such a systematic approach would be impossible to implement for 1996–2002 data, largely because the additional equations become quite cumbersome for a meaningful estimation. For example, based on a VAR OLS, a "proper" wage determination equation would have to include more than a dozen variables, most of these monthly dummies. This would be in addition, at the very least, to a separate equation for the CPI, even abstracting from the possible separate equations for the exchange rate and services prices. Since in addition the seasonality pattern starkly differs across the CPI, services prices, and wages, the sensible VAR results may be impossible to obtain on the basis of some six years of observations of very noisy data. It rather makes more sense to impose exante theoretical structure, than continue with the unrestricted model.

B. Dynamics

On the basis of the weak exogeneity assumption, we may thus proceed toward a single-equation model of inflation. To model dynamics and the long-run jointly, an unrestricted autoregressive distributed lag of the prices and the "markup" variables for the period between 1996(2) and 2002(9) was estimated, initially with six successive lags and monthly dummies. Furthermore, three dummy variables (*Bread, Sep2000*, and *Fdefl*) were added in the unrestricted regression to capture the idiosyncratic developments in food prices, respectively: (i) the upward adjustment of administered bread price increases in 1999/2000 (1 for November and December 1999 and January, February, May, and June 2000—months that were characterized by substantial increases in bread and staples prices, which were due to decisions of local governments); (ii) a sharp increase in meat prices in September 2000 that does not conform to any reasonable seasonal pattern (1 for September 2000) and (iii) food price deflation of 2002 in gross violation of the usual seasonal pattern (1 for February and June 2002—months in which sharp food price deflation of more than 1 percent was registered). No significant loss of information is found in restricting the model to two lags and one seasonal dummy (for July).

Table 1 in the Appendix presents the autoregressive distributed lag for the consumer price index. The solved long-run solution takes the form (standard errors are in parentheses):

$$p = \underset{(1.65)}{6.80} + \underset{(0.07)}{0.25}w + \underset{(0.10)}{0.42}p_{serv} + \underset{(0.05)}{0.22}e - \underset{(0.06)}{0.16}S6 - \underset{(0.10)}{0.22}Fdefl + \underset{(0.11)}{0.23}Sep2000 + \underset{(0.05)}{0.10}Bread \quad (10)$$

All the coefficients are significant and have intuitive signs, and the results are not much different from those of Table 6. The sum of the coefficients is reasonably close to unity, with those for wages and administrative prices slightly lower than in the VAR, and that for the exchange rate somewhat higher.

Table 2 in the Appendix shows the unrestricted error correction representation for the inflation rate. The number of lags in the ADL could be reduced from five to two without significant loss of information. The table makes clear that the second lag is also insignificant. The long-run homogeneity of the markup seems broadly satisfied, as the sum of the coefficients on lagged wages, exchange rate, and services prices roughly equals the negative of that for consumer prices.

A sequential reduction of a restricted error correction model embedding a long-term relationship (EC) derived in equation (10) yields the following equation for inflation, which, in addition to the terms used in the markup long term analysis, included a term for lagged broad money growth, which came out significant,[15] five monthly dummies, and the above three food dummies, all of which are highly significant.

Equation (11)

$$Dp_t = \underset{(0.04)}{0.33}Dp_{t-1} + \underset{(0.02)}{0.26}D\,p_t^{serv} + \underset{(0.01)}{0.13}De_t + \underset{(0.01)}{0.04}Dw_t - \underset{(0.01)}{0.03}EC_{t-1} + \underset{(0.02)}{0.11}Dm_{t-1} + 5Seas + 3Dum$$

DW = 2.12; R-squared=0.955.

The more detailed equation results are presented in Table 3 of the Appendix.

C. Model Evaluation

The restricted inflation model has a sensible economic interpretation. All signs are intuitive and the coefficients are significant and plausible. The fairly high coefficient on lagged inflation (about one-third) indicates substantial inertia found in other transition inflation studies (Nikolic (2000)). The short-run coefficients are plausible, with the elasticity for the services prices and the exchange rate roughly conforming to the respective shares of services and traded goods in the consumption basket. The short-term elasticity on wages is much smaller, possibly indicating a longer time lag for the pass-through, or interaction with seasonals or other short-term variables. For each of the three variables cointegrated with prices, the long-run coefficients

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[15] The effects of broad money growth term here are viewed as short term, since the cointegration analysis of the previous section did not accept the inclusion of money into a long-run cointegrating relationship. Other potential I(0) explanatory variables (i.e., interest rates) have been tried, but came out highly insignificant.

exceed the short-term ones. The coefficient on lagged money growth is highly significant. While it is not very large, it is comparable to that for the exchange rate.

The "food dummies" have the right signs and are significant. It might have been preferable to model the food market directly, but this has not been generally possible for lack of relevant monthly data for much of the sample period. The chosen representation is reasonably parsimonious, as two of the three dummies are "composite," while all dummies firmly relate to hard stylized facts.
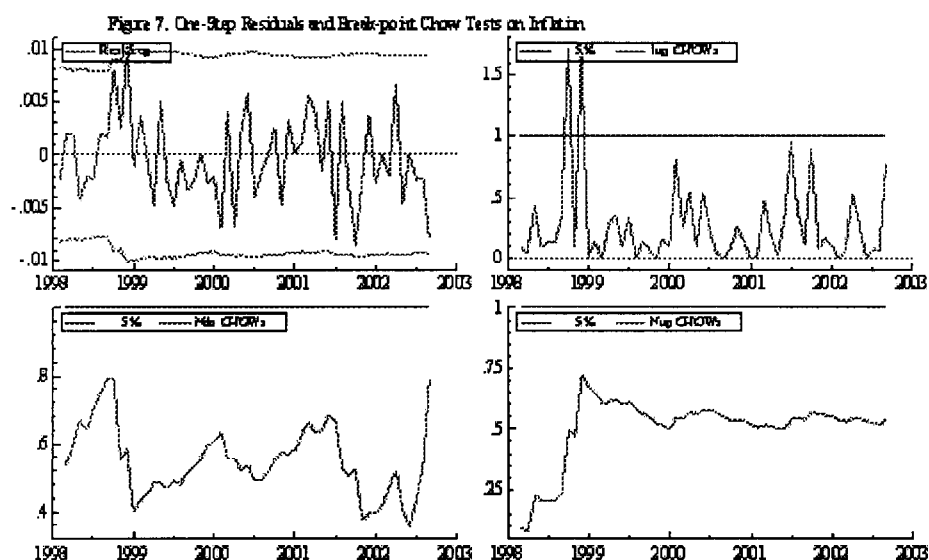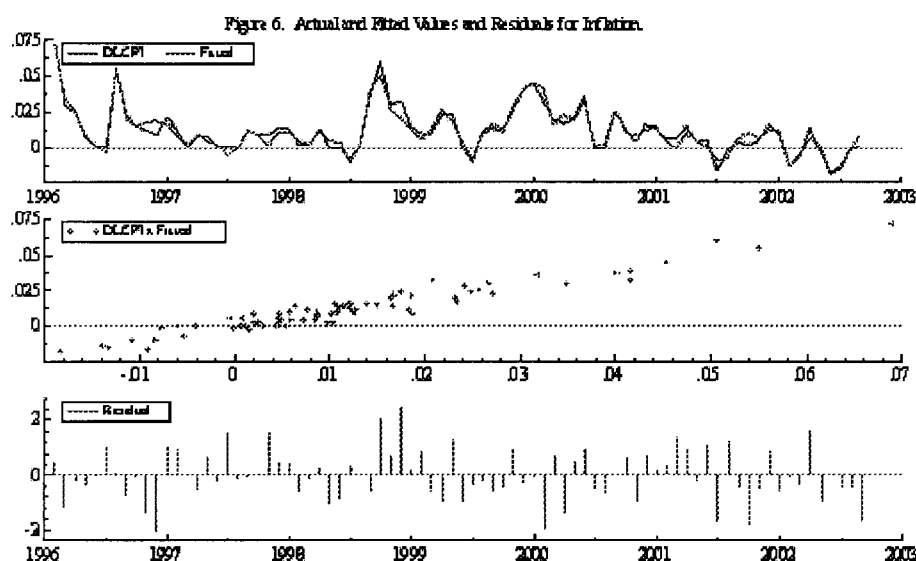
As in De Brower and Ericsson (1998), the adjustment coefficient on the markup is negative and small, indicating slow adjustment toward the long-run equilibrium. The underlying idea is that permanent increases in costs increase the CPI in the long run, whereas temporary increases in costs have small short-run effects. Interestingly, the short-run elasticity of the markup would be more than twice as high if money growth was excluded from the regression. However, this deletion is not statistically accepted (and thus is not incorporated into the final model), although the overall statistical properties of the reduced model remain fairly good.

These statistical properties of the model in equation (11) can be assessed with the help of a battery of tests. Table 11 lists the diagnostic statistics for the restricted single-equation model that test against various alternative hypotheses – residual autocorrelation (AR), autoregressive conditional heteroscedasticity (ARCH), skewness and excess kurtosis (Normality), heteroscedasticity quadratic in regressors (Xi^2) and RESET (RESET), i.e., Ramsey's statistic. The model appears well-specified, with no rejections from these tests (the probability of nonrejection is indicated in square brackets). Figure 6 confirms that the residuals are normally distributed, homoscedastic, and serially uncorrelated.

Table 11. Diagnostic Statistics for Equation (11)

AR 1- 5 F(5, 61)	=	0.6027	[0.6980]
ARCH 5 F(5, 56)	=	0.44963	[0.8118]
Normality Chi^2(2)	=	0.3451	[0.8415]
Xi^2 F(20, 45)	=	0.70443	[0.8005]
RESET F(1, 65)	=	1.0422	[0.3111]

In addition to the full-sample tests, the recursive least squares (RLS) estimation was performed over subsamples of the data. Figures 6–8 summarize output from these tests. In particular, the one-step residuals remain within two standard errors with the exception of one observation in late 1998. The standard error varies fairly little over time. The recursive "one-step" and "break-point" Chow statistics (Figure 7) are generally not significant at 5 percent levels, except one-off 5 percent levels for a couple of observations, again in late 1998. That particular time was a period of heavy exchange restrictions introduced in the aftermath of the August 1998 currency crisis, and there was substantial evidence both of the adverse velocity shock and of a higher exchange rate pass-through, as the importers quoted prices using the depreciated (by 10–20 percent) informal market rate, at which they could actually purchase foreign exchange for hryvnia. Indeed, Figure 6 shows that actual inflation exceeded fitted inflation in late 1998.

Figure 6. Actual and Fitted Values and Residuals for Inflation

Figure 7. One-Step Residuals and Break-point Chow Tests on Inflation

The recursive estimates of the individual coefficients Figure 8 generally experienced a shock at the time of the 1998 currency crisis, but afterwards vary very little, both numerically and statistically, pointing to the relative stability of the inflationary process in Ukraine during this period, despite the ongoing structural changes. Interestingly, the figure seems to confirm the anecdotal evidence on velocity and exchange rate pass-through, as the recursive estimates of both inflation inertia (a proxy for adverse velocity shocks in times of rising inflation) and the exchange rate jumped in late 1998, slightly overshooting their long-run levels.

Equation (11) could also be used for forecasting. Figure 9 plots the actual, fitted, and forecast values of inflation, focusing on the forecast period and including two standard-error bands for

Figure 8. Recursive Coefficient Estimates for Equation (11)
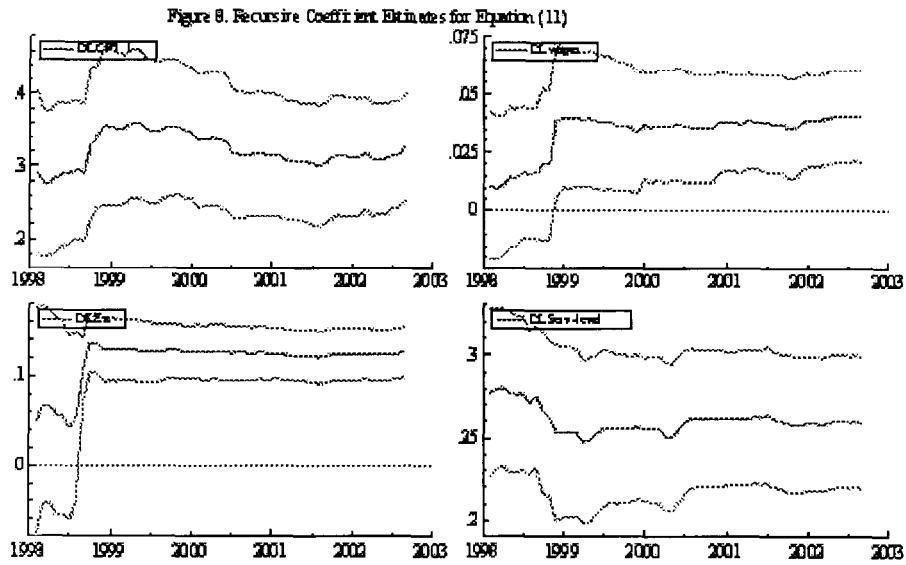
the forecasts. Estimation is through February 2002, and the realized values of the right-handside variables are used to construct forecasts, which track the recent changes in inflation fairly well. While future forecasting may also be useful with this model, there are important caveats. First, the paths of the right-handside variables in equation (11) should be given, with no feedback from the CPI on those variables. Second, this assumes both structural stability of the inflation equation and robustness to changes in the processes for wages, the exchange rate, services prices, and money growth. Finally, exante forecasts would require modeling and forecasting the right-handside variables, and these forecasts are themselves uncertain.

Figure 9. Actual, Fitted, and Forecast values of the Consumer Inflation Rate, with 2 Forecast Standard Errors
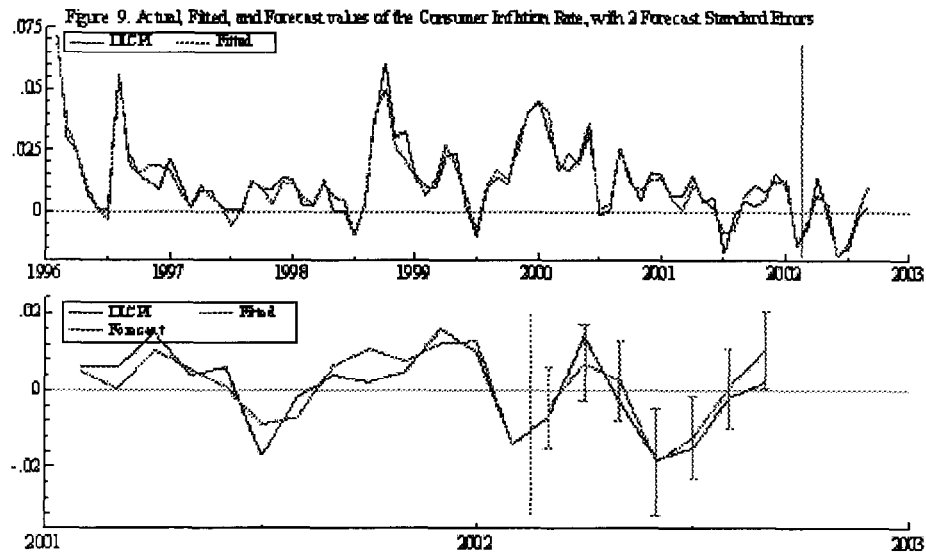
VI. THE POLICY DEBATE

The recent inflation performance raised several specific policy issues, which have been widely debated in Ukraine since 2002. First, it has been argued that further inflation stabilization effort is unwarranted, in part because inflation levels "should not be artificially reduced below a certain threshold corresponding to the economy's level of development." According to Korablin (2002), a reduction of Ukraine's inflation to low single digits could be counter productive, as it would either require "enormous effort," or "stifle growth." Second, the low inflation has raised questions as to the level of key interest rates. For example, while inflation fell by about 20 percentage points between 2000 and 2001, interest rates on bank loans declined by only a few percentage points, and, at over 20 percent, appeared extremely high in real terms in 2001–2002. Third, the huge scope for remonetization in 2000–2002 not only surprised some observers and policymakers, but also led many of them to believe that such remonetization could continue for a long period of time. In this respect, a debate ensued whether a further monetary easing would be safe to undertake.

The model in the previous section offers some, albeit partial, insights into these policy issues. On the first argument, while the model cannot assess a claim on an adverse impact of low inflation on the economy, it does not find any "threshold effects" based on Ukraine's inflation performance to date, which would indicate that achieving a progressively lower rate would require a major effort. On the contrary, the established high value of the inertia term in equation (11) prompts a pattern of virtuous/vicious circles: the higher (lower) inflation makes it more difficult (easy) to control it. This argument is perhaps more persuasive because the model controls for the fact that actual inflation in 2000-2002 was high/low for certain "exogenous" reasons (i.e., the food dummies).

Second, the level of bank interest rates depends not only on inflation, but on factors that determine risk, including protection of property rights and growth/profitability in the economy. As to inflation, it would matter in the forward-looking sense, but so will other variables, including the expected exchange rate and growth. The model and the data indicate that the expected inflation did not come down in 2002 as much as actual inflation, partly because of food deflation (likely considered temporary, and which the model controls through a dummy), and partly because of lingering expectation of tariff adjustments during 2002 that in the end never materialized. Equation (11) projects that, without the food deflation dummy, actual inflation would have been about 3 percentage points higher in 2002. This may, albeit to a rather small extent, explain why bank rates did not follow actual inflation.

The issue of scope for remonetization was extensively discussed in Section 4(IV.)(B.). While considerable uncertainty remains, the model indicates a nonnegligible (0.11) and statistically significant short-term linkage between lagged broad money growth and inflation, even under a clear period of significant remonetization. A sharp monetary expansion would feed into inflation directly, but also through the inflation inertia term, and perhaps eventually through

other monetary transmission pathways like wages and the exchange rate, if not interest rates.[16] The long-run monetary framework and the output gap seem important here, but their assessment remains elusive for Ukraine. One clear factor that appears to have helped Ukraine's inflation performance over 2000–2002, and in particular which could partly offset the impact of significant money growth was a significant tightening of the government's fiscal position, which has yet to be incorporated into a consistent analytical framework.

VII. CONCLUDING REMARKS

This paper has taken a broad approach toward assessing the evolution of prices in Ukraine, initially for 1993–2002, but especially focusing on the period of moderate-to-low inflation in 1996–2002. First, the integrating properties of the main economic time series were tested and unrestricted cointegrating relationships were examined with the Johansen procedure. After sequential reductions and assumptions, a single equation for inflation was derived for 1996–2002. The resulting model embeds a version of a long-term markup of prices over wages, the exchange rate, and administrative prices, as well as short-term factors and dummy variables. Statistically, the model performs quite well against the background of obvious problems and limitations with the time span and data quality. An added indirect confirmation of the basic results is a broad similarity of long-run coefficients obtained with unrestricted VAR and the single equation model.

The main contribution of this paper is the application of the modern cointegration analysis— now common for inflation studies in industrialized and developing countries—to the circumstances of a less-advanced transition economy experiencing remonetization. While VAR cointegration methods have been applied to some more advanced transition countries (see Kuijs (2002) for Slovakia), accounting for Ukraine's strongly seasonal data required additional comparisons between alternative theoretical models and special assumptions based on institutional arguments for a sequential reduction to a single equation. Interestingly, the high full-sample variation of actual inflation data in Ukraine has helped illuminate issues that may not be detected in lower-inflation countries.

In particular, the strong contrast between cointegration of money and prices for the whole sample (or early subsamples) and the rejection of the cointegration for 1995/6–2002, seems due to the high variation in the data during 1993–95, which apparently overpowered the obvious seasonality "noise" implied in monthly data. Interestingly, neither the data imperfections nor the extensive remonetization have invalidated this cointegration result. Against this background, the cointegration detected for the markup model during 1995/6–2002 may result from a somewhat better control for the adhoc "noise" in this model, as it offers a consistent and logical framework for including the non-modelable services part of the CPI as one of the explanatory variables.

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[16] Anecdotal evidence suggests that the role of interest rates in Ukraine's economy has been fairly small, and this has been indirectly confirmed as this variable turned out to be insignificant.

The full monetary transmission mechanism cannot be evaluated within the markup model, although this result may not be inconsistent with an underlying long-term link between money and prices.

A number of further caveats apply. Data imperfections, combined with the necessity to use monthly data and the pronounced seasonality in many series, obscure some of the relationships that could be established using a longer timeframe or better data. During 1995/6–2002, the basic "markup" results appear robust to the choice of the sample period, although the long- and short-run elasticities should not be taken literally, both due to data imperfections and the ongoing structural change in Ukraine's economy. A certain seasonality pattern also explains the statistical rejection of the intuitive assumption of weak exogeneity. Finally, some of the simplifying assumptions have likely impeded a more precise estimation of elasticities.[17] However, accounting for these latter factors seems impractical, both because of data uncertainties and due to loss in the degrees of freedom.

Despite its limitations, the markup model generates many useful conclusions for Ukraine. First, it produces a consistent and statistically well-behaved model of long- and short-term determinants of inflation. Among the former, wages and the exchange rate can hardly be independent of money, as monetary transmission mechanisms may include the exchange rate and nominal wage formation, and through those variables affect inflation (De Brouwer and Ericsson (1998)). Second, while not directly assessing the long-run monetary transmission mechanism, the model points to a significant short-term link from money to prices under an alternative, logically coherent, framework. Third, the model partly confirms the conclusions of other studies that relative price variability plays an important role in the inflation process in transition. The high long-run elasticity on services prices in the linearly-homogeneous long-run equation is consistent with this variability and with the sustained real appreciation observed in the stabilized transition economies. Finally, the model throws light on actual experience and corrects several policy assertions, pointing to a reasonably mainstream nature of the inflation process, which has been somewhat counterintuitive lately.

As the quality of data improves and a longer timeframe becomes available, future research should concentrate, inter alia, on: (i) deeper explanations of the monetary transmission process, including money demand, which should come out in sharper relief once the rate of current remonetization subsides; (ii) an allowance for the output gap; (iii) an explicit model of the food market; and (iv) incorporation of the fiscal position into the formal analysis.

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[17] The model abstracts, inter alia, from: (i) international prices; (ii) exchange rates versus currencies other than the U.S. dollar; (iii) unit labor costs (instead proxied by wages); and (iv) basic determinants of food prices (proxied by dummies).

Table 1. Autoregressive Distributed Lag Representation for LCPI, 1996–2002.

Analysis of lag structure

Lag	0	1	2	Sum
LCPI	-1	1.34	-0.432	-0.0896
StdErr	0	0.0751	0.0769	0.0324
Constant	0.61	0	0	0.61
StdErr	0.267	0	0	0.267
Lwages	0.0194	0.013	-0.0101	0.0222
StdErr	0.0121	0.012	0.0114	0.00909
LEx	0.107	-0.076	-0.0115	0.0193
StdErr	0.0199	0.0366	0.0248	0.00893
LServ-level	0.271	-0.306	0.0728	0.0373
StdErr	0.0293	0.0481	0.0342	0.0164
Seasonal_6	-0.0142	0	0	-0.0142
StdErr	0.00245	0	0	0.00245
Bread	0.00893	0	0	0.00893
StdErr	0.00296	0	0	0.00296
Sep2000	0.0207	0	0	0.0207
StdErr	0.00609	0	0	0.00609
Fdefl	-0.0196	0	0	-0.0196
StdErr	0.0046	0	0	0.0046

Table 2. The Unrestricted Error Correction Representation for Consumer Inflation

Analysis of lag structure

Lag	0	1	2	Sum
DLCPI	-1	0.221	-0.0156	-0.795
StdErr	0	0.107	0.104	0.134
Constant	0.372	0	0	0.372
StdErr	0.334	0	0	0.334
DLwages	0.107	0.0883	0.046	0.241
StdErr	0.0507	0.0598	0.0476	0.12
DLEx	0.113	0.0245	0.0309	0.168
StdErr	0.0195	0.025	0.0263	0.0377
DLServ-level	0.26	-0.00909	-0.00562	0.245
StdErr	0.0289	0.0403	0.0393	0.0599
LCPI	0	-0.0531	0	-0.0531
StdErr	0	0.0464	0	0.0464
Lwages	0	0.012	0	0.012
StdErr	0	0.0124	0	0.0124
LEx	0	0.00963	0	0.00963
StdErr	0	0.0112	0	0.0112
LServ-level	0	0.0218	0	0.0218
StdErr	0	0.022	0	0.022
Fdefl	-0.0166	0	0	-0.0166
StdErr	0.00455	0	0	0.00455
Sep2000	0.0174	0	0	0.0174
StdErr	0.00648	0	0	0.00648
Bread	0.0105	0	0	0.0105
StdErr	0.00302	0	0	0.00302
+11 Seasonals				

Table 3. Restricted Error Correction Model for Consumer Inflation

EQ(8) Modelling DLCPI by OLS (using dataukrinf1016.in7)
The present sample is: 1996 (2) to 2002 (9)

Variable	Coefficient	Std Error	t-value	t-prob	PartR^2
DLCPI_1	0.32838	0.036614	8.969	0.0000	0.5493
DLwages	0.040970	0.0099807	4.105	0.0001	0.2034
DLEx	0.12826	0.014017	9.150	0.0000	0.5592
DLServ-level	0.25871	0.019989	12.942	0.0000	0.7174
DLM2_1	0.10717	0.023010	4.657	0.0000	0.2474
Sep2000	0.015756	0.0048119	3.274	0.0017	0.1398
Fdefl	-0.015634	0.0036572	-4.275	0.0001	0.2168
Bread	0.0093564	0.0021253	4.402	0.0000	0.2270
Seasonal_2	-0.0099175	0.0021083	-4.704	0.0000	0.2511
Seasonal_4	-0.0076027	0.0019458	-3.907	0.0002	0.1879
Seasonal_5	-0.0091415	0.0020750	-4.405	0.0000	0.2272
Seasonal_6	-0.019779	0.0020402	-9.695	0.0000	0.5875
Seasonal_7	-0.0066067	0.0022808	-2.897	0.0051	0.1128
ECMfinal_1	-0.025274	0.010600	-2.384	0.0200	0.0793

R^2 = 0.954931 \sigma = 0.00470885 DW = 2.12
RSS = 0.001463436965 for 14 variables and 80 observations

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