# COINTEGRATION APPROACH TO ANALYSING INFLATION IN CROATIA

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Article**
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*Abstract*

*The aim of this paper is to analyse the determinants of inflation in Croatia in the period 1994:6-2006:6. We use a cointegration approach and find that increases in wages positively influence inflation in the long-run. Furthermore, in the period from June 1994 onward, the depreciation of the currency also contributed to inflation. Money does not explain Croatian inflation. This irrelevance of the money supply is consistent with its endogeneity to exchange rate targeting, whereby the money supply is determined by developments in the foreign exchange market. The value of inflation in the previous period is also found to be significant, thus indicating some inflation inertia.*

*Key words: inflation, Croatia, cointegration*

## 1 Introduction

Investigating the determinants of inflation is very important in the light of Croatian efforts to join the European Union (EU), given that an inflation rate in line with the requirements of the Maastricht criteria is one of the prerequisites for accession. The diversity of empirical results on this topic for transition economies makes it obvious that drawing general conclusions is an impracticable task. The aim of this paper, therefore, is to investigate the determinants of Croatian inflation empirically through the cointegration approach.

Although inflation has been widely analysed, we find that there is still room for improvement in both the model and the methodology used. This paper differs from previous

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ones on the topic in the following respects: it specifies the theoretical model on which the empirical analysis of inflation determinants is based, which is usually not the case in empirical investigations, where theory is often relied upon rather loosely; vector autoregressions (VARs) are frequently used as the main empirical tool, whereas we show that if cointegrating relationships between the variables exist, the models that use VAR are, in essence, misspecified; in addition, we use the cointegration approach, which has not often been applied to Croatian inflation data. Finally, we take into account some of the specificities of the analysed period and country in our empirical investigation, thus broadening the model, in order to model the Croatian inflationary process better.

The paper is organised as follows. Section 2 reviews the literature on inflation determinants in transition economies; Section 3 presents a theoretical model of inflation determinants; Section 4 explains the choice of the variables in our model, and explains and addresses the problems that arise with the use of certain variables; Section 5 analyses the long-run relationship between inflation, exchange rate, wages, productivity and money growth via a cointegration approach, while Section 6 concludes.

## 2 Literature review

The empirical literature on inflation determinants is substantial; a large number of studies can be found for advanced as well as for transition economies (see, for example, Pujol and Griffiths, 1998; Domac and Elbirt, 1998; Haderi et al., 1999; Nikolić, 2000; Festic, 2000; Ross, 2000; Golinelli and Orsi, 2001; Kim, 2001; Brada and Kutan, 2002; Chionis et al., 2002; Vostroknutova, 2003; Maliszewski, 2003; Masso and Staehr, 2005; Siloverstovs and Bilan, 2005; etc.). However, despite significant policy implications, the determinants of Croatian inflation have received limited attention from economists.

Given the large number of the empirical studies on inflation determinants in other transition countries, this literature review focuses only on those papers that investigate Croatian inflation. There are only a small number of studies on the subject, which we present below.

Payne (2002) analysed the inflationary process in Croatia for the period January 1992 - December 1999. He estimated an augmented vector autoregressive (VAR) model of the log first-differences for the following monthly variables: broad money supply M4; retail price index; nominal net wage per employee; and nominal effective exchange rate index. His results suggest that inflation was positively influenced by wage growth and currency depreciation in the period under investigation. However, lagged values of inflation seem not to affect present inflation.

Botrić and Cota (2006) analyse sources of inflation in Croatia in the period 1998:1-2006:3. They use two approaches, the structural VAR (SVAR) estimation following Diboglu and Kutan (2005) and the unrestricted VAR model in which they replicate Payne's (2002) empirical investigation on data from a later period. The results of their estimated SVAR indicate that terms of trade and balance of payment shocks are the most important factors generating inflation. In their unrestricted VAR analysis they find, like Payne, that the exchange rate is an important determinant of inflation. Variance decompositions based on the VAR model, furthermore, indicate that there is some degree of inflation inertia in the system, which was not, as they note, found by Payne. The impact of wages seems not to be as important in this later sample, as in the earlier one, used by Payne. Money growth is found not to be important in influencing inflation.

The findings of Payne and Botic and Cota seem plausible and in line with findings for other countries in transition. However, the empirical approach (VAR) they employ is, in our opinion, inadequate for this sort of analysis. VAR analysis has been widely adopted for analysing inflationary dynamics, as it does not require any a priori assumptions regarding the exogeneity of variables in the model, and it provides a convenient means to summarise the empirical channels with respect to economic relationships. However, if the series under consideration are cointegrated, the VAR form is not the most suitable model setup, as it omits the error-correction term. This is explained in more detail in section 5.

Vizek and Broz (2007) analyse inflation in Croatia in the period 1995-2006 using the cointegration approach. They find that mark-up and excess money are the most significant variables for explaining the short-run behaviour of inflation. Furthermore, output gap, nominal effective exchange rate, import prices, interest rates and narrow money are also found to be important in their influence on inflation. Surprisingly, nominal effective exchange rate is found to affect inflation negatively, implying that kuna depreciation actually lowers inflation. The authors hypothesise that this may mean that monetary policy reacts excessively to depreciation pressures thus causing price contractions.

## 3 Theoretical model

A commonly used model for analysing inflation determinants is that developed by Bruno (1993). One of its appealing features is that it incorporates both demand-pull and cost-push ingredients. The model starts from the balance between aggregate demand and aggregate supply:

$$Y^s\left(\frac{W}{P}, \frac{P_n^* E}{P}\right) = Y^d\left(\frac{M}{P}, \frac{EP^*}{P}\right) \tag{1}$$

where $Y^s$ is aggregate supply; $Y^d$ is aggregate demand; $P$ is the price level; $W$ is the nominal wage level; $E$ is an exchange rate; $M$ is the money supply; $P_n^*$ is exogenous import price index; $P^*$ is exogenous export price index (both in foreign currency). The relationship is next log-differentiated in order to observe the relationship between the rates of change of the four nominal variables, under the assumption that the goods market balance always holds. Accordingly the following equation is obtained:

$$\pi = a_1\omega + a_2\varepsilon + a_3\mu + v\ , \tag{2}$$

where $\pi = \frac{\dot{P}}{P}$ is rate of inflation[1]; $\omega = \frac{\dot{W}}{W}$ is wage inflation; $\varepsilon = \frac{\dot{E}}{E}$ is rate of devalua-

[1] A dot represents a discrete ($P_t$-$P_{t-1}$) or instantaneous time change (dP/dt).

tion; $\mu = \frac{\dot{M}}{M}$ is rate of monetary expansion; and $v$ represents supply and demand shocks.

Let us briefly explain in what way the above variables affect inflation. Wage increases can affect inflation in two ways. First, wage increases in excess of productivity gains can put direct upward pressure on prices. However, whether they act as an important cost-push factor depends on their share in the production costs. Secondly, they influence the purchasing power of customers through money expansion, thus affecting the aggregate demand and acting as a demand-pull factor.

As for the exchange rate, Kamin et al. (1998) note that this channel works through both aggregate demand and aggregate supply effects. On the demand side, exchange rate changes can have contradictory effects. The first is the relative price effect whereby domestic currency depreciation positively affects the price competitiveness of the country, increases the demand for domestic goods, which become less expensive relative to foreign goods, and thus increases aggregate demand and inflation. Secondly, changes in the exchange rate may also exert significant balance-sheet effects. Given that domestic residents in transition countries are net debtors to the rest of the world, and given that their debts are usually foreign currency debts, domestic currency depreciation may lead to a worse balance-sheet position, which may give rise to a contraction of domestic demand. Thus the balance-sheet effect works in the opposite direction to the relative price effect. As noted by Kamin et al. (1998), in small open economies with flexible exchange rates, the exchange rate channel is likely to be particularly important because it affects not only aggregate demand, but also aggregate supply. Namely, the depreciation of the domestic currency that results from loose monetary policy raises the domestic prices of imported goods, thus contributing to inflation directly[2]. In addition, the higher prices of imported inputs contract aggregate supply, reducing output and increasing inflation.

Finally, increases in money supply generally lead to increases in aggregate demand and prices. There are several transmission mechanisms through which changes in money supply affect aggregate demand. The interest rate channel explains this influence through negative effect of money growth on interest rates. A decline in interest rates positively affects consumer and investment spending, aggregate demand and output. Interest rate changes also negatively affect the value of asset prices, mainly those of bonds, equities and real estate (asset prices channel) thus affecting investment and consumption, and, consequently, aggregate demand. Finally, money growth and the resultant decline in interest rates improve the balance sheets of borrowers, and lead to more credits and investment (credit channel), which increases economic activity and inflationary pressures. Most studies, however, suggest that it is the exchange rate channel that is the strongest in transition economies (see Besimi et al. (2006)).

The theoretical model developed by Bruno (1993) prescribes the main selection of variables, but the empirical specification still requires difficult choices with respect to other (specific) variables, lag lengths, etc. The precise form in which different factors affect inflation is generally not specified by theory, and this applies especially to countries

[2] Under the assumption that the substitution between domestic and foreign goods is limited.

in transition. In our opinion, therefore, in order to model inflationary process in transition better, the main model should additionally include variables that reflect the specificities of the period and the countries under investigation. In the following sub-sections we will first explain the choice of each variable in our model, as well as difficulties connected to certain variables, and then turn to the estimation of the model.

## 4 Data and variables description

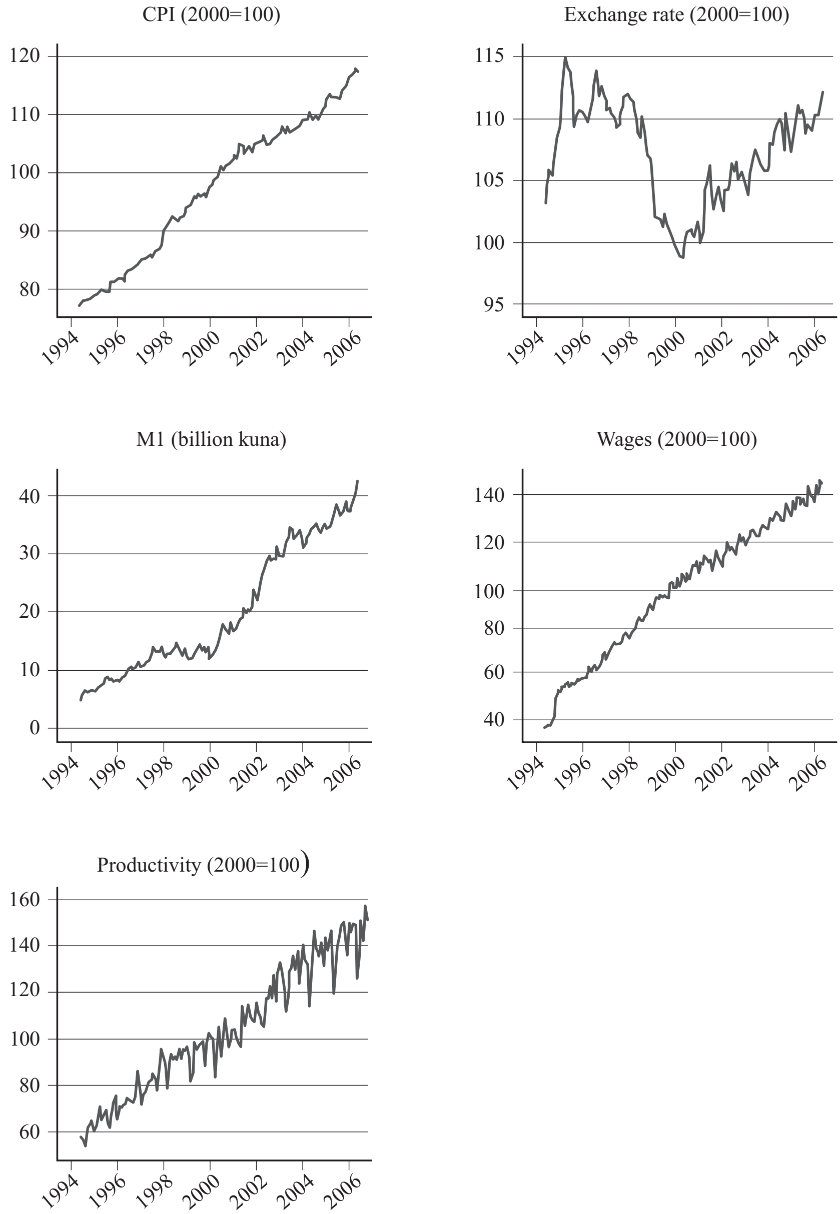
### *4.1 Core variables*

Inflation is our main variable of interest. We use the consumer price index (CPI) published by the Croatian National Bank (CNB). It should also be emphasised that inflation in Croatia was measured by the retail price index (RPI) until 1998 and by the CPI afterwards. We account for possible disparities that might occur through the inclusion of a *dummy* variable, as explained later. This variable is in further text labelled *cpi* and its logarithm *lcpi*.

As for the monetary aggregates, the CNB regularly publishes data on M1 and M4. Different studies use different monetary aggregates in order to assess their impact on inflation. Domac and Elbirt (1998) find, for example, that, in Albania, high liquid money (M1) is better at predicting the CPI than the broader definitions (M2 and M3). They explain this by the possibility that in Albania the function of money as a medium of exchange is more relevant than its function as a store of value. The same could be said for Croatia, for the level of euroisation is very high in Croatia. As argued by Billmeier and Bonato (2004), residents in Croatia maintain large proportions of their savings in foreign currency. Kraft (2003) also observes that Croats prefer foreign exchange as a store of value. Indeed, Croatia is found to display one of the highest degrees of asset substitution (holding of foreign rather than domestic money as a store of value) among transition countries (Feige, 2003). As noted by Kraft (2003), there is anecdotal evidence that foreign exchange is in use as transactions money as well, but mainly in an unofficial way since it is impossible to make payments in a store or via a bank account in foreign exchange. For these reasons we hypothesise that domestic money (kuna) serves more as a medium of exchange, while foreign currency is used as a store of value. Therefore, we decide to use the M1 monetary aggregate in our further analysis. However, as a robustness check, we will also test M4 instead of M1 (see below). This variable is in further text labelled *m1* and its logarithm *lm1*.

As mentioned previously, wages influence inflation through two channels. The first is on the supply side, through increased production costs, while the second is on the demand side, through increased demand for final goods. However, the former is true only if nominal wage increases are in excess of productivity increases. It would be desirable, therefore, to use unit labour costs (ULC) as a determinant of inflation, rather than nominal wages. ULC are given by the ratio of nominal wages per period to labour productivity. However, given that a labour productivity series is not available for Croatia, we calculate a proxy for this series by dividing industrial production with the number of persons employed in the industry. We do not attempt to calculate ULC, but we rather put both variables (nominal wages and productivity) in our model. The reason for this is that both wages and productivity (and their logs) are likely to be I(1) variables. However, their ratio

*Figure 1 depicts the above mentioned core variables and their dynamics*



(ULC) is likely to be stationary. Moreover, in each case, their growth rates are likely to be I(0). Cointegration analysis models relationships between I(1) variables; therefore, it makes sense to include the two variables separately. In addition, productivity may, through the Balassa-Samuelson effect, be a direct determinant of inflation. Namely, an increase in productivity in the tradables sector would lead to an increase in the domestic relative price of non-tradables, resulting in an increase in the price index. It should be noted, though, that Funda et al. (2007) find the influence of Balassa-Samuelson effect on inflation in Croatia not to have been statistically significant in the period 1998 to 2006. The data on average monthly nominal net wages we will be using is from the IMF's IFS online database. Wages are in further text labelled *w* and its logarithm *lw.* It is expressed as an index number with a base in 2000. The data on monthly industrial employment and production is taken from the Croatian Bureau of Statistics, and is also expressed as an index number with a base in 2000. Productivity is in further text labelled *prod* and its logarithm *lprod.*

In order to capture the 'pass-through' from exchange rate movements to the price index it is necessary to determine the exchange rate that matters the most for Croatia's CPI. Therefore, we need to use a weighted-average measure of the relevant exchange rate, i.e. an effective exchange rate. This is because the kuna might depreciate against one currency, but appreciate against some other. The effective exchange rate would summarise the effects of the combined influence of bilateral exchange rate movements and explain better their consequences for the Croatian economy. We use the nominal effective exchange rate (NEER) from the IFS database, expressed as an index with a base in 2000. The NEER is defined by the IFS in such a manner that an increase in the index reflects an appreciation (i.e. an indirect quote is used). This variable is in further text labelled *er* and its logarithm *ler.*

The variables described so far are the core variables prescribed by the theoretical model. However, it is important additionally to take into account some of the specific features of transition economies. Therefore, additional variables will be used.

### *4.2 Administered and agricultural price/supply shock dummies*

In addition to above-explained core variables in our model we include the following dummy variables:

- *D1* - reflects an increase in prices of agricultural products and a large increase in prices of telecommunication services that took place in May 1998. It has the value 1 in May 1998, and zero otherwise;
- *D2* - reflects an increase in prices of agricultural products and a high increase in prices of telecommunication services that took place in May 1999. It has the value 1 in May 1999, and zero otherwise;
- *D3* - takes account of an increase in prices of oil products. It has the value 1 in August 1998, and zero otherwise;
- *D4* - accounts for an increase in prices of telecommunication services. It has the value 1 in August 2001, and zero otherwise;
- *D5* - represents an increase in prices of food and prices of oil derivatives. It has the value 1 in February 2005, and zero otherwise.

- Value added tax (VAT) dummy (*D_VAT*) - that takes the value of 1 in January 1998 (and 0 otherwise) and represents the introduction of VAT[3].

As Mohanty and Klau (2001) point out, a large part of the movement in inflation seems to come from two major components of the price index, namely, food and oil prices. Oil prices affect inflation in countries in transition more than in industrial countries because of a more energy-intensive production structure and more energy-consuming technologies in production, transportation and heating. Therefore, it is important to include this variable as it represents a cost-push shock (Arratibel et al., 2002). Accordingly we additionally include the price of oil as exogenous variable (we label its logarithm *loil*). For this we use world oil price index from IFS (series 00176AADZF). However, petroleum product prices in Croatia were administratively regulated until 2001. Starting from 2001, these prices were determined according to an equation that takes account of crude oil world prices and the kuna/dollar exchange rate. Therefore, only after 2001 did the prices of oil products in Croatia start to reflect the changes in oil prices in the world market. We thus put an activation/deactivation dummy in front of the variable *loil* (this dummy takes the value of 0 in the period before 2001, and value of 1 afterwards), thus creating variable *loil_shift.*

## 5 Empirical analysis of inflationary process in Croatia

VAR analysis has been widely adopted for analysing inflationary dynamics because it is easily applied and requires no a priori assumptions regarding the exogeneity of variables in the model. When data is I(1) VAR is usually estimated in differences. However, estimating $y_t$ as a VAR in first differences is inappropriate if $y_t$ has an error-correction representation. Namely, if the series under consideration are cointegrated, we must include error-correction terms to allow these series to catch up with one another. The omission of the expression $\Pi y_{t-1}$,, that captures long-run adjustment, leads to standard omitted variable bias (Hess and Schweitzer, 2000). Therefore, the papers that use VAR in the case when cointegration between the variables in the model exists, are, in essence, misspecified. This is the reason we will first test for cointegration between the variables in our model, and later on use it as our main empirical tool. In addition, we want to investigate whether long-run relationships between variables exist and for this reason use cointegration. An important characteristic of this method is its ability to detect long-run relationships.

Our sample starts in June 1994 (we use monthly data), i.e. it is the period after the structural break (Stabilisation Programme in October 1993). June 1994 is chosen for several reasons: inflation in Croatia became more stable from this period onward, the official data on monetary aggregates starts in this month and we wanted to include as many observations as possible. Therefore, we will use the sample 1994:6-2006:6 in our analysis.

It is likely that monthly data would incorporate some sort of seasonal variation (because of the tourist season, Christmas shopping, harvest season etc.). Some authors (see

[3] Given that the measure of inflation in Croatia from January 1992 to December 1997 was the RPI and the CPI afterwards, we also want to include a dummy that takes the value of 1 in 1998:1 and 0 otherwise, in order to capture possible effects that arise from the change in inflation measure. However, the definition of this variable is the same as D_VAT. Therefore we include only D_VAT, although if this dummy turns out to be significant we will not be able to tell whether it is due to the introduction of VAT or to the change in price index.

for example Masso and Staehr, 2005, Enders, 2003: 196, Lutkepohl et al., 2004) use monthly dummies to account for seasonal effects. However, if we were to use this approach, we would have to introduce an additional 11 (out of 12, the 12th being the reference dummy) centred dummies in our regression, which would decrease the number of degrees of freedom. Since our sample is of relatively moderate size this is not the best option, as it means we would lose one whole year of observations. Therefore, we seasonally adjust the data[4].

*Table 1 DF-GLS unit root tests for the levels of the variables*

| **Variable** | | **Criterion** | | |
|---|---|---|---|---|
| | | **Ng-Perron** | **SC** | **MAIC** |
| Cpi | *No. of lags* | 12 | 12 | 12 |
| | *test statistics* | 1.982 | 1.982 | 1.982 |
| Lcpi | *No. of lags* | 12 | 12 | 12 |
| | *test statistics* | 1.421 | 1.421 | 1.421 |
| W | *No. of lags* | 12 | 2 | 12 |
| | *test statistics* | 0.673 | 0.927 | 0.673 |
| Lw | *No. of lags* | 12 | 3 | 7 |
| | *test statistics* | 0.622 | 0.288 | 0.233 |
| M1 | *No. of lags* | 12 | 1 | 6 |
| | *test statistics* | 1.661 | 0.674 | 0.972 |
| Lm1 | *No. of lags* | 12 | 1 | 1 |
| | *test statistics* | 2.552* | 1.276 | 1.276 |
| Er | *No. of lags* | 7 | 1 | 7 |
| | *test statistics* | 1.586 | 2.041 | 1.586 |
| Ler | *No. of lags* | 12 | 1 | 1 |
| | *test statistics* | 2.189 | 2.028 | 2.028 |
| Prod | *No. of lags* | 11 | 2 | 11 |
| | *test statistics* | 1.506 | 3.866*** | 1.506 |
| Lprod | *No. of lags* | 8 | 2 | 9 |
| | *test statistics* | 0.382 | 1.514 | 0.549 |

****, ** and * stand for the 1, 5 and 10 percent levels of significance, respectively*

*Source: Author's calculations*

We start by testing for unit roots. For this we use the Dickey-Fuller Generalised Least Squares (DF-GLS) procedure (implemented in Stata 9). The DF-GLS performs the modified Dickey-Fuller *t*-test proposed by Elliot et al. (1996). This test has significantly higher power than the previous versions of the augmented Dickey-Fuller test (StataCorp, 2004). The DF-GLS procedure optimises the power of the ADF test by generalised least squares detrending (Harris and Sollis, 2003: 57-58; StataCorp, 2004: 66). The conclusions from

[4] There are, however, downsizes of seasonal adjustment. As noted by Lutkepohl et al. (2004: 151), seasonal adjustment is an operation applied to univariate series individually and it may distort the relation between variables in a multivariate setting. Harris and Sollis (2003: 63) note that "the filters used to adjust for seasonal patterns often distort the underlying properties of the data".

the Dickey-Fuller test, even though similar, are not as strong as those from DF-GLS. The DF-GLS reports the results of three different criteria for choosing the lag length, namely, Ng-Perron, Schwartz Criterion (SC) and Modified Akaike Information Criterion (MAIC). Sometimes these results differ and in those cases we should use our own judgement to facilitate conclusions about the stationarity of the variable in question. We test the level and the logarithm of each variable, include a trend and use seasonally adjusted data. The results are given in Table 1.

The DF-GLS results indicate that we cannot reject the null of a unit root for most of the variables. The exceptions are the logarithm of money supply, where we can reject the null according to Ng-Perron criterion at ten percent, and the level of productivity, where we can reject the null according to Schwartz criterion at one percent.

In addition, in order to verify whether the variables in our model are really I(1) processes, the common next step is to test the first differences of the variables. The results (not reported) indicate that differences seem to be I(0); and hence the variables in levels are I(1). There is not enough evidence to accept two unit roots for either of the variables. The time span at hand is not likely to be long enough to identify two unit roots. In addition, the consequences of over-differencing are as serious as those of under-differencing, and for this reason also we are not inclined towards concluding that there is more than one unit root (Harris and Sollis, 2003: 58; Banerjee et al. 1994: 139 point out the loss of information; more generally, according to Hendry, 1995: 21, economic times series 'are more erratic and less systematic the more times they are differenced'[5]). We proceed using the logarithms of the variables and treating them as being integrated of order one.

*Table 2 Testing for cointegrating rank*

| Endogenous variables | Deterministic terms[a] | Lags | Johansen trace test ($H_0$: $r=r_0$) $r_0$ | LR | pval | 90%, | 95%, | 99% | Suggested number of cointegrating vectors |
|---|---|---|---|---|---|---|---|---|---|
| 1 | 2 | 3 | 4 | | | | | | 5 |
| lcpi, ler, lm1, lw, lprod | constant, trend | 3 | 0 | 147.56 | 0.0000 | 84.27 | 88.55 | 96.97 | 2 |
| | | | 1 | 82.83 | 0.0004 | 60.00 | 63.66 | 70.91 | |
| | | | 2 | 41.14 | 0.0732 | 39.73 | 42.77 | 48.87 | |
| | | | 3 | 11.04 | 0.8683 | 23.32 | 25.73 | 30.67 | |
| | | | 4 | 3.11 | 0.8538 | 10.68 | 12.45 | 16.22 | |
| lcpi, ler, lm1, lw | constant, trend | 3 | 0 | 102.43 | 0.0000 | 60.00 | 63.66 | 70.91 | 1 |
| | | | 1 | 40.33 | 0.0877 | 39.73 | 42.77 | 48.87 | |
| | | | 2 | 10.49 | 0.8964 | 23.32 | 25.73 | 30.67 | |
| | | | 3 | 2.59 | 0.9077 | 10.68 | 12.45 | 16.22 | |

[a] *We also tested the cointegrating rank upon including additional impulse dummies that refer to specific events. The LR and p-values change only slightly upon inclusion of these dummies, and the cauclusions remain the same.*

*Source: Author's calculations*

[5] One reason for the adverse effects of over-differencing noted by Hendry is that differencing an I(0) series introduces a moving average structure into the residuals and, hence, autocorrelation.

Another practical issue that should be addressed before testing for cointegrating rank refers to determining the order of the VAR (the lag length). The number of lags suggested by information criteria is given in Table 2 in column 3, together with the cointegrating rank results, columns 4 and 5. We test the rank with productivity included and excluded from the model. This is because wages can, as argued previously, influence inflation in two ways. On the supply side they have an impact on inflation only if wage increases are higher than productivity increases. Therefore, in this case it seems important to include productivity. On the demand side, on the other hand, wages influence inflation through increased purchasing power. Therefore productivity is not critical for wages to influence inflation in this case. In addition, the theoretical model does not include productivity directly. The results are presented in Table 2.

The results suggest that, when productivity is excluded there is 1 cointegrating vector (CV), and when it is included there are 2 CVs. It should be noted that, even though we use monthly data, we do realise that the ability of the cointegration test to detect cointegration depends on the relationship between total sample length and the length of the long run rather than on the mere number of observations. Hakkio and Rush (1991) note that analyses that reject cointegration usually do so as a result of the very low power of these tests when relatively small samples are used (irrespective of the number of observations). On the other hand, they note that this lack of power suggests that accepting cointegration may be a fairly strong conclusion. The next step in our analysis is to search for an adequate model for the five-dimensional system of interest (*lcpi, ler, lm1, lprod, lw*).

In vector error correction models (VECMs) the Johansen reduced rank maximum likelihood (ML) approach has been the dominant method for estimating cointegration parameters. However, Bruggemann et al. (2005) find that the Johansen ML estimator (MLE) has to be used carefully in applied work, since it can produce extremely distorted and unreliable estimates in small samples. Another problem is that such a model can pass all the usual diagnostic tests; hence these checks would not help in detecting distorted estimates. Therefore the authors suggest using a simple Generalised Least Squares (GLS) estimator, which has, in some respects, better small sample properties than MLE. The simple GLS does not produce similarly outlying estimates and is, in other aspects, very similar to MLE. We estimate a VECM with cointegrating rank 2 and 2 lagged differences using the GLS estimator[6]. In addition we include deterministic variables (*D1-D5, loil_shift* and *D_VAT*) that take account of some of the specific events that took place in Croatia, as explained above, and a trend restricted to the cointegrating relation. The model we use is given below:

$$\Delta Y_t = \nu + \alpha\left[\beta' Y_{t-1} + \tau(t-1)\right] + \sum_{i=1}^{k-1} \Gamma_i \Delta Y_{t-i} + \phi D_t + \varepsilon_t \tag{3}$$

where $Y_t$ *(m*1)* vector of m different (endogenous) time-series; $\nu$ is a *(m*1)* vector of constants; $\alpha$ is a *(m*r)* matrix of loading coefficients, r being the number of cointegrating

[6] It should be noted that our results are not robust to the empirical methodology, for were we to use the dominant empirical strategy for this sort of analysis, the Johansen's ML estimator, the results would differ in terms of sizes, signs and significances. Bruggemann et al. (2005) also find that the results obtained by the ML and GLS estimator differ markedly. It is precisely because of the Johansen criticism voiced in the paper by Bruggemann et al. that we choose to present the (more meaningful) GLS results.

vectors; β' is a *(r*m)* matrix of cointegrating coefficients; $Y_{t-1}$ is a *(m*1)* vector of endogenous time-series lagged once; $\tau$ is a *(r*1)* vector of coefficients on the time-trend *(t-1)*, which is a *(1*r)* vector and restricted to the cointegrating vector; $\Gamma_t$ is a *(m*m)* matrix of coefficients on each differenced lag (k being the number of lags) of the endogenous variables; D is a vector of exogenous dummy variables; $\phi$ is a matrix of coefficients on D. Finally, $\varepsilon_t$ is *(m*1)* vector of white noise disturbances.

The ordering of the variables is important at this point since the normalisation is such as to set the first (*r*r*)[7] block of the cointegration matrix to the identity matrix. This means that we do not get the estimated coefficients, or their *t*-ratios, for the first two variables that we list. We decide to put *lcpi* as the first variable, since it is this variable in which we are primarily interested, and *lprod* as the second variable. It was shown in testing for cointegrating rank in a bivariate setting (not reported) that productivity is possibly cointegrated with all other variables. In addition, we try different orderings (always keeping *lcpi* as the first variable) and we notice that productivity is non-significant in the first cointegrating vector and has a very small coefficient; hence we lose nothing if we do not obtain the coefficient on *lprod* in the first cointegrating vector. The two cointegration relations we obtain from the GLS estimation are[8]:

$$\begin{aligned} lcpi_t &= \underset{(4{,}755)}{0{,}139}\, lw_t - \underset{(-4{,}960)}{0{,}315}\, ler_t + \underset{(6{,}473)}{0{,}002}\, t + ec_{1,t} \\ lprod_t &= \underset{(8{,}651)}{0{,}297}\, lw_t + \underset{(7{,}852)}{0{,}216}\, lm1_t + \underset{(2{,}789)}{0{,}001}\, t + ec_{2,t} \end{aligned} \tag{4}$$

where estimated *t*-ratios are given in parenthesis, and $ec_t$ denotes the deviations from the estimated cointegration relation. The coefficient on the *lm1* in the first CV, and *ler* in the second CV were not significantly different from zero at conventional levels, so we excluded them from the equation. The cointegrating vectors indicate that there are two long-run equilibrium relationships. In the first one, *lcpi* is positively correlated with wages and negatively with the exchange rate, while in the second one productivity is positively correlated with wages and money supply. In both cases, the sign and size of coefficients are as expected. Namely, as discussed before, an increase in wages increases prices either through increasing production costs, or through increased purchasing power and increased aggregate demand. In our model an increase in wages by one percent leads to an increase in the price level of 0.139 percent (*ceteris paribus*). Increase in the exchange rate (since it is defined in an indirect quote) means currency appreciation, which negatively influences prices. More precisely, a one percent kuna depreciation induces a 0.315 percent increase in prices (*ceteris paribus*). As for the second cointegrating vector, wages influence productivity positively perhaps through the efficiency wage reasoning, whereby a rise in wages increases workers' motivation, induces less shirking and more effort. An increase in wages by one percent brings an increase in productivity level of 0.297 percent, *ceteris paribus*. As for the money supply it could influence productivity positively if its increase leads to an increase in aggregate demand and output, thus increasing productivity. In the

[7] *r* being the rank, i.e. the number of cointegrating vectors.
[8] All calculations are done in JMulti.

second CV, money (M1) growth of one percent induces a productivity growth of 0.216 percent, *ceteris paribus*. The significant linear trend in both equations may proxy technological progress or the effect of restructuring of the economy towards higher productivity sectors.

We next turn to the analysis of the loading coefficients, i.e. the coefficients on the cointegrating vectors. These coefficients measure at what rate per period one of the endogenous variables adjusts to correct a temporary disequilibrium in the cointegrating vector and so moves the variables back towards their long-run equilibrium relationship. Table 3 reports these adjustment coefficients together with their associated *t*-statistics. For reasons of space, the five error-correction models are not reported in full. We firstly note that CV1 enters all equations, but the second one. CV2, on the other hand, enters the productivity, exchange rate and money supply equations.

*Table 3 Adjustment coefficients and their associated t-statistics (from the error-correction models)*

| **Cointegrating vector** | **Dependent variable in the error correction model (ECM)** | | | | |
|---|---|---|---|---|---|
| | **lcpi** | **lprod** | **lw** | **ler** | **lm1** |
| CV1 | -0.329*** | | 0.356*** | -0.049*** | -0.093* |
| | (-6.510) | | (3.348) | (-2.749) | (-1.756) |
| CV2 | | -0.712*** | | -0.073*** | -0.144* |
| | | (-5.943) | | (-2.737) | (-1.800) |

*t-ratios in parenthesis; ***, ** and * stand for the 1, 5 and 10 percent levels of significance, respectively*

*Source: Author's calculations*

The loading coefficient for the first equation (with *lcpi* as the dependent variable) is rather high (-0.329). It means that if the price level (*lcpi*) is temporarily above (below) its long-run equilibrium level, then inflation falls (rises) by, approximately, 33 percent per month until equilibrium is restored. This implies a period of only six to seven months for 90 percent of adjustment to take place. The adjustment coefficient in the ECM with wages growth as the dependent variable indicates that if the price level is above the equilibrium level, wages rise in such a way that 36 percent of adjustment is accomplished in each month, in order to restore the equilibrium. The adjustment coefficient in the ECM with percentage changes in the exchange rate (indirect quote) as the dependent variable is -0.049, meaning that if the past price level has been too high (low) in relation to long-run equilibrium, then adjustment towards equilibrium will be achieved through exchange rate decrease (increase) (and thus currency depreciation (appreciation)). This adjustment will be approximately 5 percent monthly, meaning that it would take three and a half years for 90 percent of adjustment to take place in this manner alone. Finally, the adjustment coefficient in the ECM with *lm1* growth as the dependent variable is -0.093, meaning that if the past price level has been too high in relation to long-run equilibrium, then subsequently money supply will adjust to restore long-run equilibrium 9 percent each month. This finding is not in line with expectations, but since it is not highly significant (only at the ten percent) we do not read too much into it. It should be emphasised that the above adjustments assume that there were no changes in other variables at the same time. As for the loading coefficients on the second cointegrating vector, the number -0.712 suggests that if the productivity level is above the equilibrium level, productivity decreases so as to restore equilibrium within three months (i.e. approximately 71 percent in each successive month). The adjustment coefficient in the ECM with percentage changes in the exchange rate as the dependent variable indicates that if the productivity level is above the equilibrium level, the exchange rate will decrease in such a way that 7.3 percent of adjustment is accomplished in each month, in order to restore equilibrium. This means that, in the absence of changes in the other variables, it would take the exchange rate two and a half years for 90 percent of adjustment. Finally, the adjustment coefficient in the ECM with *lm1* growth as the dependent variable is -0.144, meaning that if the past productivity level has been too high in relation to long-run equilibrium, then subsequently money supply will adjust to restore long-run equilibrium 14 percent each month. This finding, as in the previous case, is not in line with expectations, but since it is not highly significant (only at the ten percent) we do not read too much into it.

Finally, in each error correction equation most short-run determinants and deterministic variables proved not to be significant, so we do not comment on them for each equation. Let us just note that in the inflation equation, the one we are mostly interested in, the first lag of inflation is significant, thus suggesting some inflation inertia. Interestingly, none of the other endogenous variables (wages, exchange rate and money supply) is a significant short-term determinant of inflation in our model. Among the set of deterministic variables, *D_VAT* is significant and positive, thus indicating that the introduction of VAT increased inflation (although, as mentioned before, a measure of inflation in Croatia from January 1992 to December 1997 was the RPI and the CPI afterwards, so this significance might actually reflect the effect of this change in the price index, rather than the introduction of VAT, since the two changes were simultaneous). In addition, *loil_shift* is also significant among the short-term variables. This suggests that oil prices also positively influenced inflation after 2001, when prices of oil products in Croatia started to reflect the changes in oil prices in the world market.

## 6 Robustness checks

Since neither theory nor empirical work on this topic give clear answers as to what variables should be included in the regression and in which way, we undertake some robustness checks.

Some authors (Payne (2002) and Botrić and Cota (2006)) use, as a measure of money growth variable, the broadest available measure - M4 instead of M1. We, therefore, repeat the whole procedure already explained before, only this time using *lm4* (logarithm of M4) as one of the endogenous variables. The first CV does not change significantly upon making this change since the *lm4*, as *lm1* before, is statistically insignificant, and therefore excluded from the regression. The coefficients and *t*-statistics on the remaining two variables in the first CV (wages and exchange rate) are very similar to those when *lm1* was used. There is a change in the second CV, however, for here only *lm4* seems to be significant, while both wages and exchange rate are insignificant at conventional levels. As for the loading coefficients, they indicate that, as before, CV2 does not enter the inflation equation ( which we are most interested in), but this time CV1 does enter the productivity equation.

We also test whether a different approach to including the oil variable influences our results. First we include only *loil*, without the activation/deactivation dummy in front of it. This does not have a significant influence on the sizes, signs or significances of our CVs. CV2 in this setting once again does not enter the inflation equation. This time, however, it does not enter the money supply equation either. Interestingly, when *loil_shift* was included, it was, as a deterministic variable, significant in all equations except the money supply equation, whereas *loil* is only significant in inflation and productivity equation. There are additional differences with respect to short-run determinants. Namely, the *loil* variable now seems to influence inflation negatively, thus suggesting that an increase in the prices of oil decreases inflation. This makes no sense, and leads us to the conclusion that the proper way to treat this variable is to include a dummy variable in front of it, as we did before. In this way we take into account the fact that oil prices in Croatia started reflecting world oil prices only after 2001. Finally, including only the mentioned activation/deactivation dummy (starting in January 2001) instead of the oil variable makes no significant changes in the results (compared with the ones in Equation 4).

## 7 Concluding remarks

Results from cointegration and error-correction modelling identified some possible sources of Croatian inflation. The analysis reveals that there is a long-run relationship between inflation, the exchange rate and wages, in the post-Stabilisation period, but not with the money supply. Increases in wages positively influence inflation in the long-run. In addition, the depreciation of the currency has also contributed to inflation in the period from June 1994 onward. Money does not explain Croatian inflation, implying possibly that monetary policy has been dedicated to other targets. This is, thus, consistent with the endogeneity of money supply to exchange rate targeting, whereby money supply is determined by developments in the foreign exchange market.

The significance and importance of the nominal exchange rate as a factor in the inflationary process is comparable to findings in Brada and Kutan (2002) for the Czech Republic, Hungary, and Poland; Haderi, et al. (1999) for Albania and Ross (2000) for Slovenia. The significance and importance of wages as a determinant of inflation is comparable to the findings of Festic (2000) for Slovenia. As for Croatia, our results are comparable to those of Payne (2002) since we all find that inflation was positively influenced by wage growth and currency depreciation in the period under investigation. However, in his analysis lagged values of inflation were not important in affecting inflation whereas our results suggest that there was some inflation inertia present. Botric and Cota (2006) also find that the exchange rate is an important determinant of inflation. They, like us, also find evidence of inflation inertia. However, they conclude that wages are not as important a factor as they were in the earlier (Payne's) sample. This finding is attributed to the fact that wages were more important at the beginning of the 1990s, when wage indexation was often explicitly stated in the collective bargaining, than in later years, when they did not play such an important role. Our analysis, on the contrary, reveals that wages are still an important factor in determining inflation. Vizek and Broz (2007) also find nominal effective exchange to be a long-run determinant of inflation, together with the mark-up, excess money, and the output gap. Their findings differ from ours, in that, firstly, we find the theoretically correct sign on the exchange rate, which indicates that currency depreciation increases prices, and secondly, unlike them, we find that money is not a significant determinant of inflation. In our sample this is true irrespective of what monetary aggregate is used, M1 or M4. Our finding that inflation in Croatia is not strictly related to monetary aggregates growth is found by most authors who investigate Croatian inflation.

The vector error correction model suggests that it takes a very short time (around seven months) for changes in inflation to return the price level to its long-run equilibrium after it has been driven away from it. The same is true for wages, *ceteris paribus*. It takes exchange rate adjustment approximately three and a half years to return the price level to its equilibrium after a disequilibrating shock. The short-run dynamics of inflation are affected by the previous-period inflation, thus suggesting a degree of inflation inertia in the system. Other endogenous variables have no short-term effect on inflation. Among the set of deterministic variables, oil prices (after 2001) proved to be an important determinant of inflation in the short-run.

As a final comment, let us just note that the last year included in our sample is 2006. Currently the whole world is experiencing structural changes which might result in the breaking-up of the previously found long-run relationships. Therefore any future research, that would include these changes in the sample, will have to be aware of these other mechanisms and it will be necessary to take account of structural breaks. The literature on cointegration in the presence of structural breaks has been growing in the past few years, after Johansen, Mosconi and Nielsen (2000) generalised the cointegration analysis in a multivariate setting developed by Johansen (1988, 1991), to the case where structural breaks exist at known points in time (see, in addition, Trenkler et al., 2006 and Trenkler, 2002). We leave this sort of analysis for future research.

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