Investigating Croatian Inflation through the Cointegration with Structural Break Approach

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Abstract: *This paper analyses the inflationary process in Croatia during the period 1992-2011, using a cointegration with structural break approach. Our results indicate that there is a long-run relationship between inflation, exchange rate, unit labour costs and money growth. Currency depreciation and unit labour costs are found to influence inflation positively, and money supply negatively. We argue that the latter occurs because exchange rate targeting policy in Croatia results in a situation where endogenous money moves in the direction opposite to the exchange rate, so as to keep the exchange rate fixed. We, furthermore provide some evidence that money supply need not mean risks to inflation in the presence of declining money velocity.*

Keywords: *Inflation; Croatia; Cointegration; Transition*

JEL Classification: *C32; E52; P24*

Article History
Submitted: 15 July 2014
Resubmitted: 25 September 2014
Resubmitted: 4 November 2014
Accepted: 10 November 2014

http://dx.doi.org/10.14706/JEOCSS11517

Introduction

Relatively low and stable inflation has been a characteristic of Croatian economy for almost twenty years. There is a rare agreement among the economists about the necessity of achieving price stability at a low level and it is widely accepted that inflationary environment is detrimental to growth, employment and even happiness. Investigating the inflation process in any country is important for several reasons: firstly, low inflation improves resource allocation and financial stability; secondly, low inflation rate (the one in line with the Maastricht criterion) is one of the prerequisites for joining the Euro Area and thirdly, in order to be able to keep inflation at the desired level central banks have to be familiar with the inflation generation process. Given the importance of inflation and its relationship with other variables in the economy, one of the main aims of this paper is to analyse the inflationary process in Croatia in the longest available period: 1992-2011.

Croatian inflation has not been investigated sufficiently, and the whole period for which the data exists has never been analysed. Namely, the early transition period is often excluded from empirical analyses mainly due to the fact Croatia, like most other transition countries, had only just established an independent bureau of statistics in those years. Therefore, the data from this period is often considered to be unreliable; alternatively it is unobtainable. Furthermore, this was a turbulent period and stabilisation programmes, adopted in most countries at the beginning of transition, represented a clear structural break in a number of data series. Since the procedures that account for structural breaks are computationally more complex, these periods are often excluded from the analyses and structural breaks are ignored.

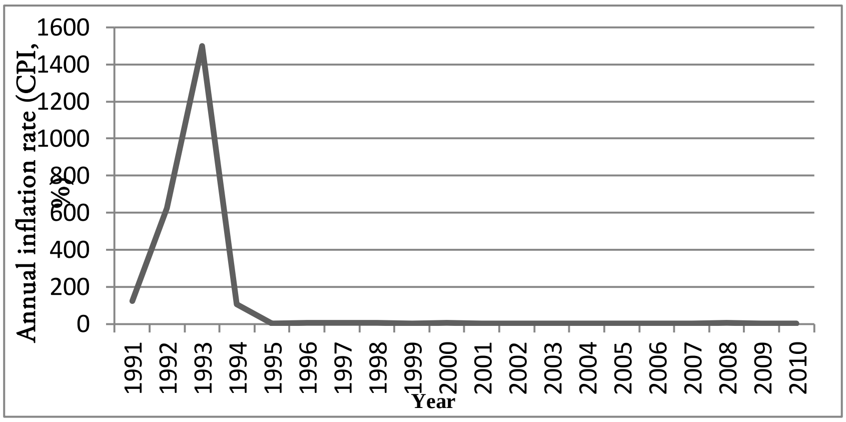
In spite of these problems we believe that it is important to account for the early period in the analysis, as it offers important explanations of the current inflation generating process in Croatia (as explained in more detail in Section 2). Namely, the Stabilisation Programme adopted in 1993 brought hyperinflation in Croatia to a halt using the exchange rate anchor, and the public in Croatia still perceives the association between the exchange rate movements and the inflation rate as to follow the same direction. Furthermore, the use of the pre-Stabilisation data in combination with the latest available data permits, for the first time, an assessment of Croatian inflation in the long-run (20 years). Finally, we use a cointegration in the presence of a structural break approach as a methodology that fits the data for Croatia in the period under investigation best. This methodology is relatively novel, and has not been used before for analysing inflation, to the best of our knowledge.

The paper is organised as follows: Section 2 gives an overview of the inflationary dynamics in Croatia since the 1990s, given that the roots of today's monetary policy in Croatia can be found in the early transition period. Section 3 reviews the literature on inflation determinants in transition economies focusing particularly on the (scarce) literature investigating Croatian inflation. Section 4 gives a theoretical model of inflation, which is based on the three hypothesis of inflation determination: excess money supply, foreign inflation and cost-push inflation. Section 5 discusses methodological issues. More precisely, data issues are first examined, since the inclusion of the early period warrants this; unit roots are investigated next, using the approach that accounts for structural breaks, and finally, cointegration with structural breaks methodology is applied to the data to assess the long-run relationships between the variables in the model. Section 6 discusses the obtained results at length and offers explanations for, what can at first be interpreted as, somewhat atypical findings. Section 7 undertakes a number of robustness checks, while Section 8 assesses whether conventional conclusions about the sources of Croatian inflation still prevail.

A History of Croatian Inflation

Croatia has been very successful at restraining inflation in the early 1990s as can be seen from Figure 1.

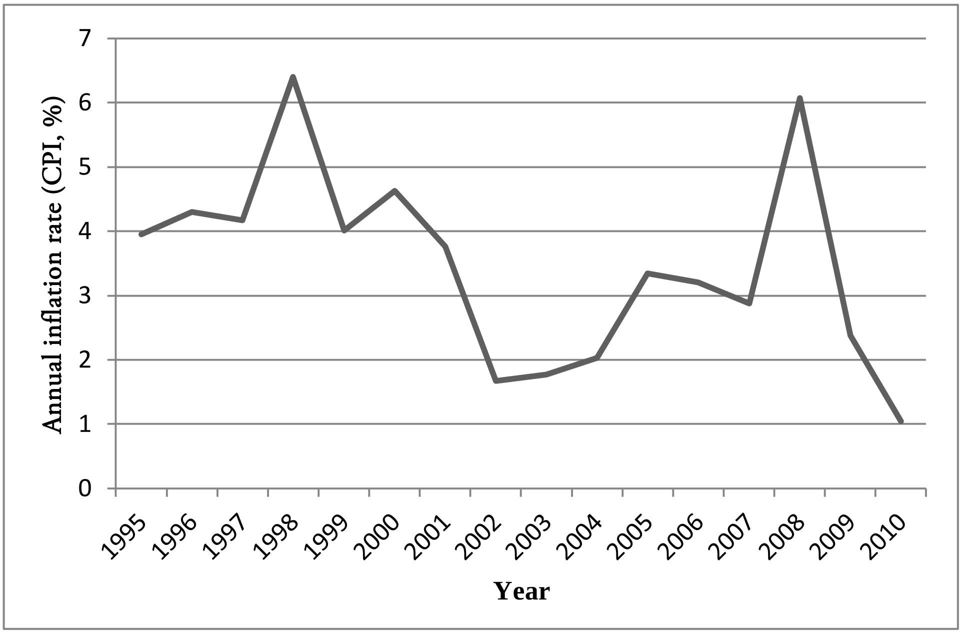
Figure 1. Annual Inflation (CPI based) in Croatia in 1990-2010



Source: International Financial Statistics

The outburst of hyperinflation at the beginning of Croatian transition, when annual inflation reached levels above 1500 percent, can be attributed to several factors. The initial outbreak of inflation was due to price liberalisation, which led to a one-time adjustment of previously repressed prices, and monetary overhang and fiscal deficit inherited from the previous regime (ex-Yugoslavia). The fiscal deficit additionally expanded due to war financing, inability to borrow money in international markets and a lack of international reserves, as well as quasi-fiscal deficits, in the form of government subsidies. This led to more money creation and further fuelled inflation, thus contributing to inflation persistence. Wage pressures (through indexation) as well as relative price adjustments (through cost-recovery hypothesis and the Balassa-Samuelson effect) also contributed to inflation persistence. As can be seen from Figure 1, Stabilisation Programme adopted in October 1993 managed to lower inflation to a single-digit number, and it was kept at relatively low levels ever since (Figure 2). This is considered one of the major economic achievements of the transition period.

Figure 2. Annual Inflation (CPI based) in Croatia in 1995-2010



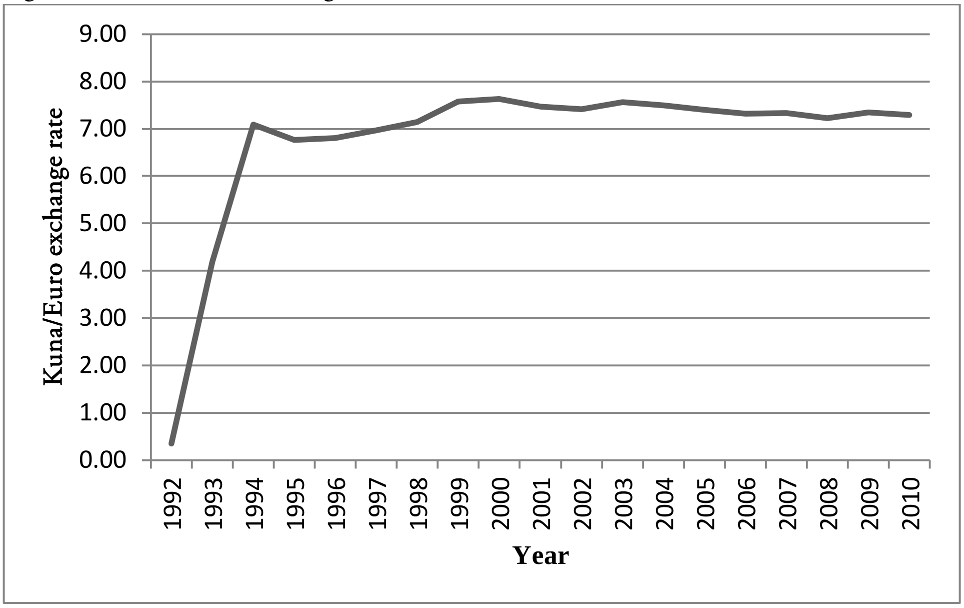
Source: International Financial Statistics

The degree of currency and asset substitution in Croatia was already high at the beginning of 1990s as a result of inflationary history in 1970s and 1980s when Croatia was still part of Yugoslavia. Indexation to exchange rate was omnipresent; even the communal services used indexation. Croatian citizens had substantial holdings of foreign exchange and depressed domestic currency incomes. Inflationary expectations were closely tied to exchange rate depreciation. Furthermore, an important factor was previous negative experience with high inflation levels, so people were very sensitive to changes in the exchange rate and as a consequence, reacted to these changes.

The Stabilisation Programme adopted in October 1993 was exchange rate based. The key element of the Programme was the introduction of the current account convertibility. This enabled citizens to convert domestic currency into foreign currency (and vice versa) at any time (Kraft, 2003). Croatian citizens and companies were in need of local currency, caused by domestic money supply tightening in months prior to the Programme. Current account convertibility enabled them to achieve liquidity by exchanging foreign currencies for local cash (Kraft and Franicevic, 1997). The exchange rate was pegged to the German mark (DM) and Croatian National Bank (CNB) announced the upper intervention point of 4444 Croatian dinars (HRD, Croatian currency at that time) for 1 DM in October 1993. A new foreign exchange law was also adopted in October. Under this law banks could freely set their exchange rates (Sonje and Skreb, 1997). Right after the implementation of the Stabilisation Programme there was an extremely high growth of cash as the central bank worked within the previously set quantitative limits. Since the central bank did not allow new credits in domestic currency to the commercial banks, and since it was no longer buying foreign currency, this unexpected and high cash growth totally exhausted commercial banks' liquidity (Anusic, 1995). Commercial banks, headed by Varazdinska banka (medium-sized bank) started offering foreign currency at lower exchange rates to attract domestic currency and increase liquidity. Namely, they started asking 4200 Croatian dinars for one German mark, which was below black market rate at the start of the anti-inflation programme. Later on, as the banks tried to discourage households from exchanging their foreign currency reserves for domestic money, they lowered the exchange rate even more, and required 3800 Croatian dinars for one DM. This turned the tide of expectations and increased the trust in the Stabilisation Programme (Babic, 1998). Due to serious monetary tightening and increasing confidence in domestic currency, as well as the new exchange rate regime and new foreign currency market arrangements, dinar started appreciating (reverse currency substitution), and returned to the pre-stabilisation level of 3708.8 dinars for one DM (Nikic, 2000). Thus, the credibility of the Programme was established.

As can be seen from Figure 3, the exchange rate remained relatively stable ever since (that is, after 1994).[i]

Figure 3. Kuna/euro Exchange Rate in 1992-2010



Source: Croatian National Bank

The exchange rate regime in Croatia is a managed float, with fluctuations in the range of +/-6 percent. In cases when foreign currency supply is above demand, CNB intervenes by increasing international reserves, and in times of increased demand it sells foreign currency, thus preventing extra depreciation. The central bank, thus, by setting the exchange rate as an intermediate target, renounces of monetary policy independent from the one in the anchor country. This means that the money supply in the domestic country becomes endogenous variable determined by the money supply in the anchor country. In such situations, as noted by Egert and MacDonald (2008), exchange rate changes provide a nominal anchor for expectations, as they may signal changes in prices. Therefore, a large(r) pass-through effect can be expected for countries with an accommodative monetary policy. Interestingly, Billmeier and Bonato (2004) find that exchange rate pass-through has been low in Croatia in the period 1994-2001. Kraft (2003) also finds the level of pass-through in Croatia to be modest.

Literature Review

This section reviews the literature that deals with the determinants of inflation in transition economies and especially Croatia. In the first part we briefly summarise the main findings of various papers on this topic, while the second part consists of a more detailed assessment of selected papers. The papers selected for the second part are those that investigate determinants of Croatian inflation, since presenting all the papers assessing inflation in transition economies would be an impracticable task.

In most studies that analyse inflationary processes in transition countries money growth is found to be the main driving force behind inflation. This holds in: Albania (Haderi et al., 1999), Russia (Nikolic, 2000), Slovenia (Ross, 2000) and in Czech Republic, Hungary and Poland (Brada and Kutan, 2002). The exchange rate seems to exert a notable impact on inflation in Poland (Golinelli and Orsi, 2001), Czech Republic and Hungary (Golinelli and Orsi, 2001), Slovenia (Ross, 2000) and in the three Baltic states (Masso and Staehr, 2005). Finally, wages are found to be an important determinant of inflation in Slovenia (Festic, 2000; Ross, 2000). Next we present the papers analysing Croatian inflation in more detail. Table 1 summarises these papers and gives their main conclusions, while a more detailed analysis is given below the table.

Table 1. Literature Review

Author(s)	Period	Data	Methodological approach	Main determinants
Payne (2002)	Jan1992-Dec1999	Monthly	VAR	Wages, exchange rate
Botric and Cota (2006)	Jan1998-Mar2006	Monthly	VAR; SVAR	Wages, exchange rate
Vizek and Broz (2009)	1995Q1-2006Q2	Quarterly	Cointegration	Mark-up, excess money, exchange rate and output gap
Malesevic Perovic (2009)	Jun1994-Jun2006	Monthly	Cointegration	Wages, exchange rate
Dolezal (2011)	Jan1998-Nov2010	Monthly	Cointegration	Money, exchange rate

Payne (2002) analyses the inflationary process in Croatia for the period January 1992 - December 1999 using a vector autoregression (VAR) approach. His results suggest that inflation is positively influenced by wage growth and currency depreciation. Money and lagged values of inflation seem not to affect present inflation.

Botric and Cota (2006) analyse sources of inflation in Croatia in the period January 1998 – March 2006. They use two approaches. In the first one they estimate a structural VAR (SVAR), following Dibooglu and Kutan (2005), while in the second approach they replicate Payne's (2002) unrestricted VAR using the 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, similar to Payne, that the exchange rate and wage growth are important determinants of inflation. In addition, their results point toward a significant positive effect of monetary growth on inflation. Variance decompositions furthermore suggest that there is some degree of inflation inertia in the system, which was not, as they note, found by Payne. Botric and Cota explain their variance decompositions as indicating that the inflation process has changed, whereby wage growth has lost its strength in explaining inflation.

VAR analysis has been widely adopted for analysing inflationary dynamics. However, when cointegration relations are present in the system of variables, the VAR form is not the most suitable model setup, since the model is in this case misspecified. Stationarity and cointegration testing should, therefore, be the starting point of any such analysis. This has been done, to the best of our knowledge, in only three papers.

Vizek and Broz (2009) use the cointegration approach and analyse Croatian inflation in the period Q1 1995 – Q2 2006. They show that the short-run behaviour of inflation is best explained via mark-up and excess money. Additional sources of the short-run inflation include: output gap, nominal effective exchange rate, import prices, interest rates and narrow money. In the long-run, mark-up, excess money, nominal effective exchange rate and the output gap are found to be the key determinants of inflation. An unusual finding is that currency depreciation affects prices negatively. The authors argue that, in the Croatian 'fear of floating' context, it might mean that monetary policy reacts excessively to depreciation pressures thus causing price contractions.

Malesevic Perovic (2009) analyses the determinants of inflation in Croatia in the period June 1994 – June 2006 through a cointegration approach. She finds wage growth and currency depreciation to be the main driving forces behind inflation in the period under investigation. Money is found not to be important in explaining Croatian inflation, and this is used as a confirmation of money endogeneity to exchange rate targeting.

Finally, Dolezal (2011) uses the period January 1998 – November 2010, and in his analysis of the monetary transmission mechanism in Croatia concludes, among other things, that there exists a cointegrating relationship between the real exchange rate, money (M1) and inflation. However, the signs on all the variables in his model are the opposite of expectations (theory).

Cointegration approach, although more appropriate than VAR, has a drawback in the sense that it should be applied to long(er) samples, and the period 1995-2006 (used by Vizek and Broz, 2009), 1994-2006 (used by Malesevic Perovic, 2009) and 1998-2010 (used by Dolezal) cannot really be considered a long-run. By using the latest available data (up till September 2011) and merging them with the earliest available period (pre-June 1994) we were able to create, for the first time, a sample of 20 years of monthly observations on inflation in Croatia, which can be considered a long-run.

Theoretical Approach

A commonly used model (see, for example, Ross, 2000; Payne, 2002; Botric and Cota, 2006; Malesevic Perovic, 2009) for analysing inflation is the one developed by Bruno (1993). In this model the main determinants of inflation (π) are wage growth (w), exchange rate changes (ε) and money growth (μ), or more formally:

$$\pi = \alpha_1 w + \alpha_1 \varepsilon + \alpha_1 \mu + \upsilon \tag{1}$$

where υ represents supply and demand shocks.

Wages may influence inflation through two main channels. The first one is on the supply side, through increased production costs, while the second one is on the demand side, through increased demand for final goods. However, since the former is true only if nominal wage increases are in excess of productivity increases, it would be more desirable to use unit labour costs (ULC) as a determinant of inflation. This is what we do in our analysis.

The impact of exchange rate on inflation works through both: aggregate demand and aggregate supply side. On the aggregate demand side domestic currency depreciation can add up to inflation through its positive effect on price competitiveness of the country, thereby increasing aggregate demand and inflation. On the supply side currency depreciation raises the domestic prices of imported goods, thus contracting the aggregate supply, reducing output and increasing inflation.

Finally, increases in money supply generally lead to increases in aggregate demand and, consequently, prices. The transmission mechanisms of monetary policy are well-know, so we do not elaborate on them at this point. The issue of the impact of money on inflation is discussed in more detail in Section 6.

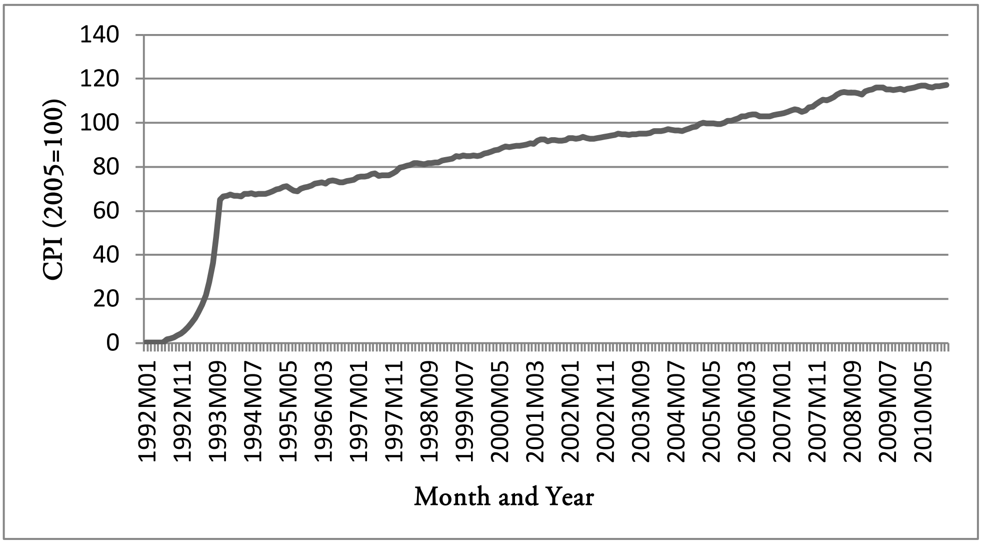
Methodological Approach

Data Issues

Changes in economic policies and economic structures that happened during transition to a market economy raise questions about the appropriate methodology to apply, as well as about the best ways to incorporate these changes in the econometric model. Some of the problems that arose are: the relatively short time span (most transition countries have the data only from 1990 onwards); seasonal adjustment of the data, which is more notable in the monthly data usually used in modelling economies in transition; measurement problems that consist of non-systematic errors; partially observed variables and systematic measurement errors (Erjavec, 2003). In this section we briefly discuss some of the mentioned problematic issues.

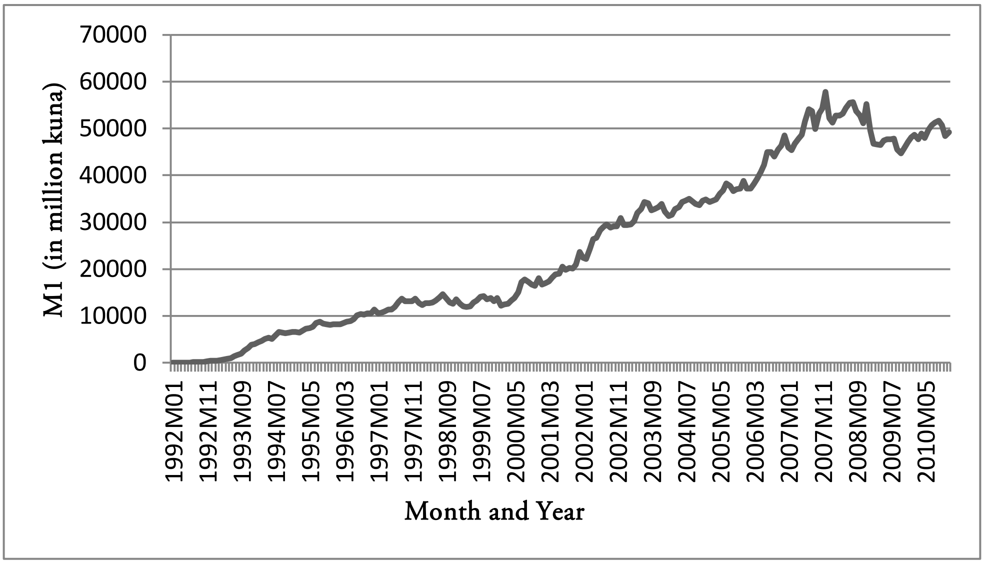
In the period under investigation a structural break occurred, when the 1993 Stabilisation Programme drastically and suddenly changed the value of all the variables in our model, as presented in Figures 4-7.

Figure 4. CPI During the Period 1992:M1-2010:M12



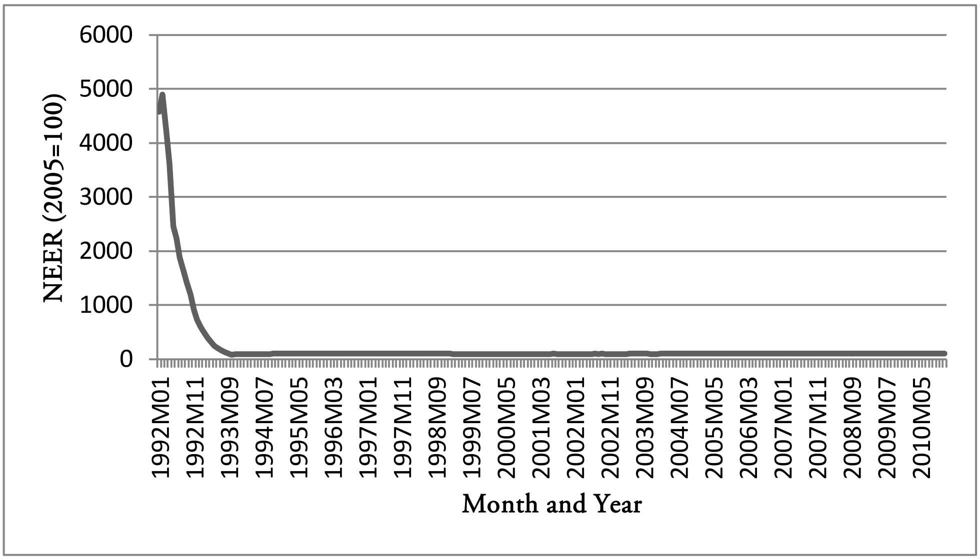
Source: International Financial Statistics and author's calculations

Figure 5. M1 during the Period 1992:M1-2010:M12



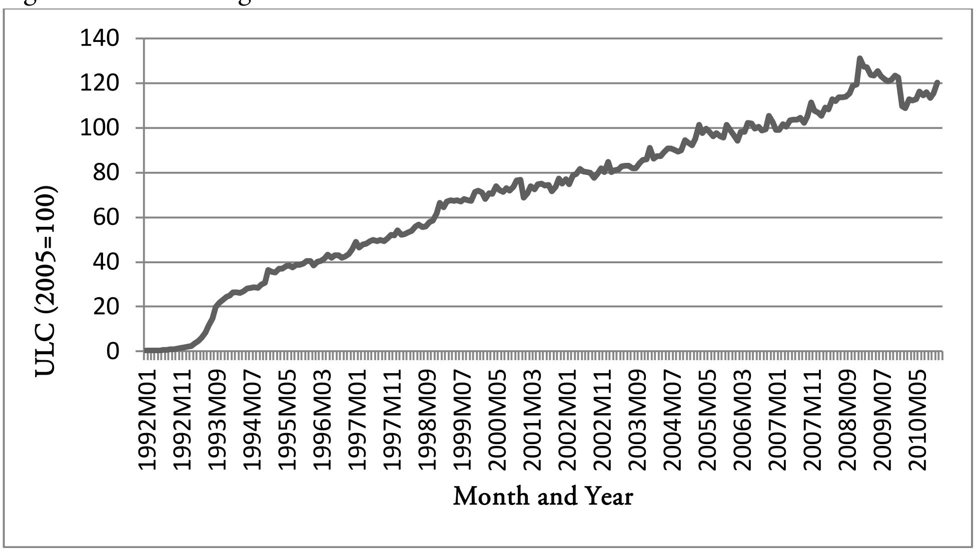
Source: International Financial Statistics and author's calculations

Figure 6. NEER during the Period 1992:M1-2010:M12



Source: International Financial Statistics and author's calculations

Figure 7. ULC during the Period 1992:M1-2010:M12



Source: International Financial Statistics and author's calculations

Structural breaks have to be accounted for in both unit root testing and in cointegration analysis. Namely, as observed by Perron (1990), conventional unit root tests may be unable to distinguish between the data generating process where random shocks exert permanent effects on the economic system and the one in which such shocks have no permanent effect but take place in the presence of a one-time permanent shift(s) in the trend function. This renders a need to develop alternative statistical procedures that can distinguish a process with a unit root from a process stationary around a trend function, which contains a one-time break (Perron, 1989).

As for cointegration, Trenkler et al. (2006) note that ignoring structural breaks can lead to seriously incorrect inference in cointegration testing. More precisely the incidence of structural changes has an influence on the conclusion about the cointegrating rank. Using an incorrect cointegrating rank can in turn lead to a wrong economic interpretation of the behaviour of the system analysed and, furthermore, it may have a negative impact on other inference tools (Trenkler, 2002). For these reasons in our empirical work we carefully account for structural breaks in both; unit root and cointegration testing.

Another problem is that of unreliable and unavailable data, especially in the first years of transition. We, however, managed to obtain the data on all the variables from 1992 onward. The data on consumer price index (CPI) is taken from the International Monetary Fund (IMF) International Financial Statistics (IFS) database. It is expressed as an index based in 2005. As for the ULC, since this series is not readily available for Croatia on monthly basis, we first calculate a proxy for labour productivity by dividing industrial production with the number of persons employed in the industry and then compute the ULC as the ratio of nominal wages per period to labour productivity. The data on monthly industrial employment and production is taken from the Croatian Bureau of Statistics. The data on monthly wages is taken from the IFS. It should be noted that the productivity measure is a rather narrow one, since it is calculated only for the industry sector. However, since the data is not available for other sectors, this is a common approach to calculating unit labour costs (see, for example, Tica and Jurcic, 2007). The nominal effective exchange rate (NEER) is taken from the IFS database, and it is expressed as an index with a base in 2005. Most problems were encountered with regards to the money supply variable. For money supply we use the M1 monetary aggregate in our empirical analysis. Although some papers suggest using a broader monetary aggregate as a determinant of inflation we believe that high liquid money is better at predicting the CPI because the function of money as a medium of exchange is more relevant than its function as a store of value in this context (which is essentially captured by M4). Croatian National Bank publishes monthly data on M1 from June 1994 onwards. The data from January 1992 to April 1994 that we use in our analysis is from an old CNB Bulletin, and it is expressed in Croatian dinars. The parity between kuna and Croatian dinar was 1:1000 at the moment of introduction of the kuna, so we convert this series to kunas. Furthermore, since we pool the data from two different periods (pre- and post-Stabilisation period) the data for May 1994 was missing. This issue was solved through imputation.

All the data in our model is monthly, seasonally adjusted and converted into logarithms. The analysed period is January 1992 – September 2011.

Unit Root Testing

The date of the structural break (T_B) in our data is known - it is the Stabilisation Programme in October 1993 (data plots in Figures 4-7 also shows this very clearly), and this should be taken into account in testing for stationarity. Hence, following Perron (1989, 1990), in testing for the existence of unit roots we run the following model:

$$\Delta y_t = \hat{\mu} + \hat{\theta} DU_t + \hat{\beta} t + \hat{\gamma} DT_t + \hat{d} DTB_t + \hat{\alpha} y_{t-1} + \sum_{i=1}^{k} \hat{c_i}\, \Delta y_{t-i} + \hat{e_t} \quad (2)$$

where y_t is a time-series that is being tested; $\hat{\mu}$ is the estimated constant; $\hat{\theta}$ is the estimated coefficient on DU; DU is a dummy variable equal to one for all periods after the structural break ($>T_B$) and zero otherwise; $\hat{\beta}$ is the estimated coefficient on the time trend, t is the time trend; $\hat{\gamma}$ is the estimated coefficient on DT; DT is a dummy variable equal to the time trend for all periods after the structural break ($>T_B$) and zero otherwise; $\hat{d}$ is the estimated coefficient on DTB; DTB is a dummy variable equal to one only in the period right after the structural break (T_B+1) and zero otherwise; $\hat{\alpha}$ is the estimated coefficient on y_{t-1}, while y_{t-1} is the first lag of y_t; $\hat{c}_i$ is the estimated coefficient on Δy_{t-i}; Δy_{t-i} stands for various ($i=1,\ldots,k$) lagged differences of the dependent variable and $\hat{e}$ is the estimated white-noise error term.

The results of unit root testing are given in Table 2.

Table 2. Perron's Unit Root Tests

VARIABLE	LEVELS		GROWTH RATES[ii]	
	t-stat	lags[iii]	t-stat	lags
CPI (no trend)	5,4242***	12	-12,0466***	12
CPI (with trend)	-1,0847	12		
M1 (no trend)	-2,3522	12	-8,5260***	11
M1 (with trend)	-2,3585	12		
NEER (no trend)	-5,2637***	12	-6,9104***	5
NEER (with trend)	-3,2281	12		
ULC (no trend)	-2,1079	12	-7,6652***	2
ULC (with trend)	-2,5455	12		

Numbers in the table are t-statistics on the coefficient α from (2). The null hypothesis is that α=0. Critical values from Perron's (1989) model C; are -4.38; -3.75 and -3.45 for 1 (***),5 (**) and 10 (*) percent, respectively.

Table 2 presents the *t*-statistics on the coefficient α in (2). These statistics are compared to critical values available in Perron (1989). A visual inspection of the data (Figures 4-7) suggested that our variables of interest are non-stationary, and this visual impression is further confirmed by the results in Table 2. Namely, the null of a unit root cannot be rejected for all the tested variables in levels. The only exceptions are the CPI and NEER without the trend. For growth rates, on the other hand, we can strongly reject the null of a unit root. Taken together, the results indicate that the variables in levels are I(1), that is, non-stationary, so our next step is to test the variables for cointegration.

Cointegration with a Structural Break

In order to empirically investigate and apply the theoretical model presented in (1) to Croatian data we use cointegration approach for two main reasons. Firstly as indicated before, VAR approach often used for investigating the sources of inflation is misspecified in the case when cointegrating relationships exist among the variables of interest. Secondly, cointegration embeds the economic notion of a long-run relationship between economic variables and our data enables us to analyse the longest possible period, which has not been analysed before.

Another issue is that a structural break occurred in the period under investigation, so the approach we actually undertook is that of cointegration in the presence of structural breaks. Johansen, Mosconi and Nielsen (2000) generalise the cointegration analysis in a multivariate setting developed by Johansen (1988, 1991), to the case where structural breaks exist at known points in time. As the authors point out, their model generalises Perron's (1989, 1990) model C (that we used for unit root testing). The major issue related to cointegration in the presence of structural breaks is that new asymptotic tables are required. They are not, as before, published or readily available. This is because the critical values depend on whether and how many trend breaks or just simple level shifts are included in the model. Furthermore, the relative break points or, to be more precise, the relative sub-sample lengths have an impact on the critical values.

The Johansen, Mosconi and Nielsen (2000) approach to cointegration in the presence of structural breaks has not been widely used by applied economists, as it is relatively new, and the computation of critical values can be difficult. In addition, it has not, to the best of our knowledge, been applied to the analysis of inflationary dynamics before. We have already stated that the incidence of structural changes has an influence on the conclusions about the cointegrating rank, which can cause wrong economic interpretation. Furthermore, Trenkler (2002) finds that ignoring level shifts leads to size distortions in such a way that the tests' size approaches zero for increasing values of the shift magnitude. For these reasons we believe it is necessary to apply the Johansen, Mosconi and Nielsen (2000) approach in our analysis.

The main logic behind the Johansen, Mosconi and Nielsen (2000) model is that an observed time series is divided into sub-samples according to the positions of structural breaks. A vector autoregression is chosen for each of the sub-samples, so that the parameters of the stochastic components are the same for all sub-samples, while the deterministic trend differs between sub-samples.

The model we use is given below:

$$\Delta Y_t = \nu + \alpha[\beta' Y_{t-1} + \tau(t-1) + \phi DT_{t-1}] + \sum_{i=1}^{k-1} \Gamma_i \Delta Y_{t-i} + \sum_{i=0}^{k-1} \gamma_i DTB_{t-i} + \eta\, DU_t + \varepsilon_t \quad (3)$$

where Y_t is a $(m*1)$ vector of m different (endogenous) time-series; ν is a $(m*1)$ vector of constants; α is a $(m*r)$ matrix of loading coefficients, r being the number of cointegrating vectors; β' is a *(r*m)* matrix of cointegrating coefficients; Y_{t-1} is a *(m*1)* vector of endogenous time-series lagged once; τ 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; ϕ is a (*r*1)* vector of coefficients on *DT*; *DT* is a (*1*r*) vector representing trend shift dummy; Γ_t is a *(m*m)* matrix of coefficients on each differenced lag (*k* being the number of lags) of the endogenous variables; γ_t is a *(m*1)* vector of coefficients on *DTB; DTB* is a (*1*m*) vector of impulse dummies, while η is a *(m*1)* vector of coefficients on *DU; DU* is a (*1*m*) vector of level shift dummies. Finally, ε_t is *(m*1)* vector of white noise disturbances. *DT*, *DTB* and *DU* are defined as in Perron's (1989, 1990) unit root tests above.

Results

The Johansen, Mosconi and Nielsen (2000) procedure tests for the rank, *r,* of the matrix $\Pi=\alpha\beta'$, where α is the matrix of adjustment coefficients, while the matrix β includes cointegration vectors. The cointegration tests check the pair of hypothesis (trace variant) $H_0(r_0)$: $rk(\Pi)=r_0$ versus $H_1(r_0)$: $rk(\Pi)>r_0$, r_0 being the tested rank of the matrix Π. The results of the Johansen trace test are given in Table 3.

Table 3. Johansen Trace Test (H_0: $r=r_0$)[iv]

r_0	p-value	90%	95%	99%
0	0.0000	70.54	74.57	82.52
1	0.0181	47.64	51.05	57.85
2	0.3023	28.61	31.37	36.99
3	0.5715	13.58	15.67	20.10

The results suggest that there exists one cointegrating vector, so we proceed using Johansen reduced rank maximum likelihood (ML) approach for estimating cointegration parameters in our vector error correction model (VECM). This procedure is based on (3), and has a trend (*t*) and trend shift (*DT*) restricted to the cointegrating relationship. The obtained cointegrating relation is given below.

$$\ln(CPI) = 0.521\ln(NEER_t) - 0.147\ln(M1_t) + 0.520\ln(ULC_t) - 0.386DT_t + 0.387t + ec_t \quad (4)$$

where ec_t denotes deviations from the estimated cointegration relation. T-statistics (not reported) indicate that all the variables in the cointegrating vector are significant at conventional levels. The results indicate that unit labour costs and exchange rate exert a positive impact on the CPI, while monetary aggregate M1 exerts a negative impact. The first two results are in line with our prior expectations. Namely, an increase in unit labour costs, which essentially means an increase in wages over and above gains in labour productivity, induces employers to pass the higher costs onto consumers in the form of higher prices, in order to protect the real value of their profits. In our model an increase in ULC by one percent leads to an increase in the price level of 0.52 percent (*ceteris paribus*). Increase in the exchange rate, that is currency depreciation, also influences prices positively, by increasing the domestic prices of imported goods. Our results suggest that a one percent kuna depreciation induces a 0.521 percent increase in prices (*ceteris paribus*). This can be interpreted as a long-run pass-through coefficient. Finally, the result on M1 variable is not in line with theory, as we would anticipate inflationary effect of excess money. Our results, however, suggest that money (M1) growth of one percent induces a fall in the prices of 0.147 percent, *ceteris paribus*. Since this finding seems unusual at first, we next investigate this relationship in more detail.

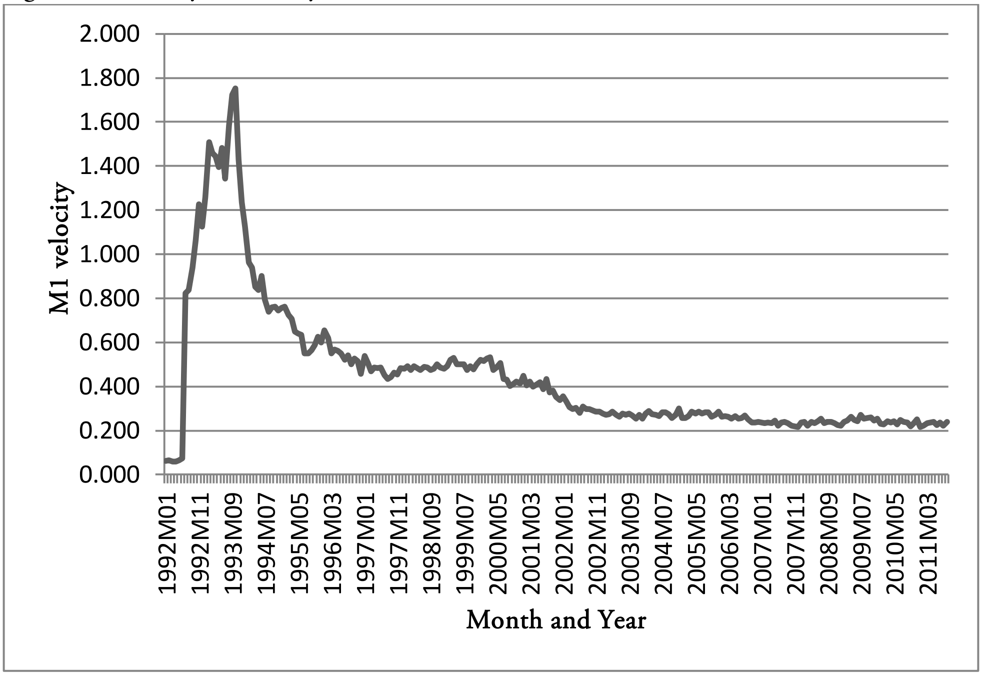
The relationship between money and prices stems from the quantity theory of money, which implies that an increase in money supply causes proportional change in the price level. The whole theory is based on the identity $MV=PY$, where M represents the stock of money; V is the velocity of money; P stands for the price level, and Y is the real income. The above identity implies that money is inflationary only under the assumption that money velocity is time invariant or constant. If, however, money velocity is time variant, then an increase in money supply may coincide with equivalent or a larger drop in velocity and, as a result, coexist with no or even negative change in nominal income (PY). This is sometimes referred to as 'velocity crowding-out of quantitative easing' (Pattanaik and Subhadhra, 2011).

A declining velocity means that the interval between economic transactions has increased, which implies the money demand is increasing (both consumers and businesses wish to hold more cash). For as long as velocity continues to decline, acceleration in the underlying rate of inflation is highly unlikely, regardless of the rate of expansion of the money stock.

To check whether this might indeed be the case with our data, we calculate money velocity from the above identity, using M1 as M, CPI as P and industrial production index (taken from the IFS database) as a proxy for real output (Y) (due to unavailability of monthly GDP data).[v] The velocity series derived in this manner is presented in Figure 8.

From Figure 8 we can clearly see a structural break in 1993, followed by a downward trend in money velocity in the period under investigation. This property of money velocity in Croatia has already been observed by Cziraky and Gillman (2006), who find this variable to be non-stationary (and descending) in the period 1994-2002.

Figure 8. Velocity of Money in Croatia in 1992:1-2011:9



Source: International Financial Statistics and author's calculations

In general, high volatility of money velocity arises as a result of high currency and asset substitution which happens as a response to economic and political instability and hyperinflation. This is a common feature of developing and transition economies and, as explained in Section 2, of Croatia also. An unstable money velocity implies that money demand is also unstable, resulting in monetary authorities not being able to rely on a dependable transmission mechanism between money supply and inflation.

As indicated in Section 2, due to exchange rate targeting, money supply in Croatia is endogenous, or, in other words, determined by the developments in the foreign exchange market. This means that an increase in the exchange rate (currency depreciation), triggers a reaction from the central bank, which sells foreign exchange and decreases money supply thereby dampening the fall in the currency value. Increase in the exchange rate leads to an increase in inflation, and since money adjusts so as to keep the exchange rate (relatively) fixed; this results in there being a negative relationship between money and prices.

In conclusion, the opposite signs on the exchange rate and money supply variables seem to reflect Croatian monetary policy quite well. Namely, given Croatia's 'fear of floating' syndrome, high level of euroisation and monetary policy focused on maintaining the exchange rate stability, the importance of the exchange rate channel in monetary transmission should come as no surprise. Interestingly, Vizek and Broz (2009) justify their finding of a negative effect of depreciation on inflation by excessive monetary reaction to depreciation pressures. Contrary, our results suggest that the central bank reaction is moderate, and that it is the impact of the exchange rate, rather than that of the money supply, that is crucial for influencing inflation. This is in line with the conclusions of Coricelli, Jacbec and Masten (2004) that an accommodative exchange rate policy is one of the main sources of inflationary pressures in accession countries.

To complete our analysis, we comment on the four associated error-correction models (ECMs). Table 4 reports the adjustment coefficients which measure the rate at which one of the endogenous variables adjusts each month to correct a temporary disequilibrium in the cointegrating vector.

Table 4. Adjustment Coefficients and their Associated *t*-statistics

Dependent variable in the ECM			
ln(CPI)	*ln(NEER)*	*ln(M1)*	*ln(ULC)*
-0.069***	0.137***	-0.001	0.394***
(-3.837)	(6.357)	(-0.010)	(5.853)

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

All the coefficients in Table 4, apart from the one in the money equation, are statistically significant, implying that the cointegrating vector enters all equations, but the third one. The loading coefficient for the first equation (with *ln(CPI)* as the dependent variable) is -0.069, suggesting that, if the price level is temporarily above (below) its long-run equilibrium level, then inflation falls (rises) by, approximately, 7 percent each month until equilibrium is restored. This implies a period of 33 months for a 90 percent of adjustment to take place. The adjustment coefficient in the ECM with percentage changes in the exchange rate as the dependent variable is 0.137, meaning that if the past price level has been too high (low) in relation to the long-run equilibrium, then the equilibrium will be achieved through currency adjustment of 13.7 percent monthly. It would, therefore, take a year and a half for the 93 percent of correction to take place in this manner alone. Finally, the adjustment coefficient in the ECM with ULC growth as the dependent variable indicates that if the price level is above the equilibrium level, ULC will adjust almost 40 percent each month in order to restore the equilibrium, suggesting that 90 percent of adjustment will be achieved after only 4 to 5 months. It should be emphasised that the above adjustments assume that there were no changes in other variables at the same time.

The results, furthermore, imply that money supply is weakly exogenous. Namely, a variable is considered to be weakly exogenous if the cointegrating relation does not enter the equation for that variable. This is in line with the aforementioned endogeneity of money supply in Croatia.

In order to preserve space, especially given that we use 10 lags in our model, the four error correction models are not reported in full. Let us just note that in each error correction equation a lot of short-run determinants are found to be statistically insignificant. Those that are significant refer mainly to the influence of inflation, exchange rate depreciation and ULC growth on inflation. This suggests that it is not only the exchange rate and ULC that play an important role in influencing current inflation, but that there exists considerable inflation inertia also. Indeed, the coefficients are found to be significant for lags 1, 5, 6, 8 and 10. Furthermore, past values of the exchange rate, ULC and inflation also exert a significant impact on the current exchange rate. Short-run coefficients are found to be mostly insignificant in their influence on money and ULC.

Robustness Checks

Robustness checks have been undertaken in order to test the validity of our results and, consequently, conclusions.

Firstly, the same model has been tested as in (4) only this time using M4 monetary aggregate for money variable, since some papers use this broader definition of the money supply (see, for example, Botric and Cota, 2006). The results do not change upon this, in terms of signs and significances of the variables in the model. We also test whether including wages instead of the ULC changes the main implications of the results, but this does not influence the findings notably. We also test the same model with twelve lags (larger number of lags as suggested by some information criteria), which, besides improving the diagnostics somewhat, has no effect on our conclusions.

In addition, given that some papers find that it is important to include oil prices as additional explanatory variable, because they represent an important cost-push shock (see for example, Mohanty and Klau, 2001; Golinelli and Orsi, 2001; Arratibel et al., 2002 and Malesevic Perovic, 2009), we, accordingly, expand the model so that it includes the influence of oil prices on inflation. Following Malesevic Perovic (2009) this issue has been approached in three ways. We firstly test whether including just world oil prices (world monthly crude oil (petroleum) prices from the IFS database have been used) as additional (exogenous) variable affects inflation and our conclusions regarding the cointegrating vector. Secondly, we use a shift dummy variable for the period after 2001, since this is the period when prices of oil in Croatia began changing in accordance with the world prices. Thirdly, we also multiply this dummy variable with the oil price variable thus including only the oil prices after 2001 in the model. In addition, given that value added tax (VAT) was introduced in Croatia in 1998 we also include additional dummy for this VAT change since it could potentially act as an additional cost-push factor. None of the above changes significantly influences our findings in terms of the signs and sizes of the coefficients in the cointegrating vector.

Given that the inclusion of the early, unstable, period in our analysis might be perturbing our results, we furthermore test only the period after June 1994, which excludes the structural break and the data of questionable quality. The results (not reported) confirm the existence of one cointegrating vector, in which money is again found to influence inflation negatively and all the other signs and significances remain the same.

Furthermore, as it was emphasised in Section 2, because of exchange rate targeting money supply in Croatia is endogenous. It might be more informative, then, to include money supply in the anchor country as a determinant of Croatia's inflation.

We test for this by including monthly broad money for Euro area, since Croatian currency was firstly tied to the German mark, and later on to the euro. The data is taken from OECD Statistics, and is compiled from the national contributions supplied by the participating countries' national central banks. Data prior to January 1999 have been converted from national currencies using the irrevocable exchange rates fixed on 31 December 1998. Upon inclusion of this monetary aggregate our results (not reported) still suggest that there is one cointegrating vector among the variables in the model, and within this vector all the variables have positive signs, that is, the signs implied by the theory. Money supply is, however, statistically insignificant in this setting.

Conclusion

Our analysis differs from other investigations of the inflationary dynamics in Croatia in several aspects. Firstly, we use a newer and longer sample that includes the last twenty years. Furthermore, we use a novel approach – cointegration with structural breaks, which enables us to include the early, pre-stabilisation, period into the analysis and account for structural breaks properly. In addition, instead of the usually-used variable in this sort of analysis – wages, we construct and use unit labour costs, which reflect the cost-push aspect of inflation better.

Our results indicate that there is a long-run relationship between inflation, exchange rate, unit labour costs and money growth. Expectedly, currency depreciation is found to add to inflation, and this pass-through coefficient is found to be quite large, contrary to the findings of Billmeier and Bonato (2004) and Kraft (2003). Unit labour costs are also identified as an important cost-push factor. Moreover, majority of adjustment towards the long-run equilibrium seems to be happening precisely through this channel. This finding contradicts Botric and Cota (2006) who conclude that, although important, wage growth has lost its strength in explaining the rates of inflation. Finally, papers dealing with Croatia typically find that money growth is not significant in its influence on inflation or that this influence is positive (see Section 2). We, on the other hand, find this influence to be statistically significant, extremely robust and negative. This negative impact of money supply is interpreted as a logical result, given the circumstances in which monetary policy in Croatia operates. Namely, exchange rate targeting, which has roots in Croatia's hyperinflationary history and the conduct of the 1993 Stabilisation Programme, results in money being endogenous in Croatia. We argue that this accommodative exchange rate policy leads to a situation where an increase in the exchange rate leads to an increase in inflation, and since money adjusts downward so as to keep the exchange rate (relatively) fixed, this results in there being a negative relationship between money and prices. Taken together, these results point towards the importance of the 'exchange rate view' at the expense of the 'money view' of the monetary transmission channel.

In addition, we argue that the transmission mechanism between money supply and inflation may further be agitated due to the shocks to money velocity, which can add significant noise to monetary analysis. Namely, money injected into the system becomes inflationary only when spent or rolled over frequently, whereas it was shown that money velocity was constantly declining in Croatia in the period under investigation. In this case growing money supply need not mean risks to inflation.

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[i] Croatian dinar (HRD) was replaced by Croatian kuna (HRK) in 1994, while German mark officially ceased to be legal tender in 2001 and was replaced by the euro. For these reasons in Figure 3 we present the kuna/euro exchange rate.

[ii] We use Perron's procedure for growth rates and not first differences, as it is unclear how and/or whether this procedure can be used with differences. As for the growth rates we just assumed that the variable was not, for example, CPI but rather inflation (that is, CPI growth). In this way we used the same procedure as we did when testing the levels of the variables.

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[iv] 10 lags were used as suggested by the Akaike information criterion.

[v] $v=p+y-m$, where $v=ln(V); p=ln(CPI); y=ln(Y)$ and $m=ln(M1)$