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Exchange Rate Pass-Through and Monetary Policy in Croatia

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Abstract

Exchange rate targeting is considered the best policy option in dollarized economies when wages and prices are indexed to the exchange rate. Croatia is a highly dollarized economy, but empirical investigation conducted in this paper shows that exchange rate pass-through has been low after stabilization. This finding, which is robust to different methodologies (VAR, cointegration), would suggest that dollarization is mostly limited to financial assets and therefore that strict exchange rate targeting may not necessarily be the best option. However, policy implications are unclear due to the endogeneity of the pass-through to the policy regime.

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I. Introduction

Monetary policy in Croatia has been very successful in reducing inflation by using the exchange rate as the nominal anchor. This policy, which could be defined as "strict exchange rate targeting", is characterized by a very low tolerance of exchange rate movements and a marked activism of the central bank on foreign currency markets. Strict exchange rate targeting has been very successful in ending hyperinflation and stabilizing the economy in the mid-1990s and delivering low inflation throughout the period following the early 1990s war. Recent changes in the legislative framework and moves towards a larger use of market-based instruments suggest that this policy may have to be revised in the future. The new Croatian National Bank (CNB) law, by clearly defining price stability as the primary objective of monetary policy, reduces the emphasis on the exchange rate. The liberalization of the capital account required as part of the European Union (EU) accession process is already testing the ability of the CNB to maintain tight control on the exchange rate. The development of financial markets provides new policy instruments and the progressive reduction of the role of reserve requirements opens up new opportunities for a more active use of open market operations by the CNB. It is thus reasonable to expect that exchange rate fluctuations will be more pronounced in the future and that the CNB will be more tolerant of these fluctuations.

A move in this direction would benefit from a better understanding of the transmission mechanism, i.e., of how exchange rate movements pass through to domestic inflation. Exchange rate appreciations or depreciations have a direct impact on inflation by changing the price of tradables expressed in domestic currency. By altering the relative price of domestic and foreign goods, the exchange rate also affects inflation indirectly through changes in economic activity. Small, open economies that are price takers in the world market may expect exchange rate movements to feed into the domestic price level, with prices at various stages of the production chain being differently affected. Ceteris paribus, import prices should move one-to-one with the exchange rate, as dictated by the law of one price. As soon as other inputs are added, the corresponding price measure should reflect the weight of imports in the production process. However, estimated pass–through can deviate from this benchmark for several reasons. Strategic pricing in foreign markets can drive the pass–through to zero.[3] The existence of menu costs associated with price adjustment ensures that mere noise in the exchange rate, as long as it is regarded as stationary and "small" relative to a threshold, is not reflected in price changes.[4] Expectations and different forms of indexation can largely affect the final result.

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[3] Consider a German car exporter to the US: if the US dollar depreciates, the law of one price predicts a rise in the USD price, ceteris paribus. In order to maintain market shares, the exporter might choose not to change the price, thereby reducing the profit margin.

[4] This threshold in general will be positively correlated with past volatility in the exchange rate.

Notwithstanding the wealth of channels through which the exchange rate can affect inflation, the available empirical evidence, mostly limited to advanced economies, points to a low and decreasing pass-through (McCarthy, 2000). The pass-through appears to be endogenous to different regimes and tends to be smaller when inflation is low, indirectly confirming the importance of expectations in the transmission mechanism (Choudri and Hakura, 2001). The evidence of a low pass-through does not seem to hold for transition economies (Ross, 1998; Kuijs, 2001), possibly reflecting lack of credibility of the monetary authorities and structural elements, like the price-taking nature of domestic firms on international markets. A number of features of the Croatian economy are likely to affect the magnitude of the pass-through. Two are worth noting. Croatia turns out to be an intermediate case, compared with other transition countries in terms of openness.[5] More importantly, Croatia is a heavily "dollarized" economy, with widespread asset substitution and some indexation of prices to the exchange rate. This, by itself, would argue for a large pass-through.

The remainder of the paper is structured as follows. Section II describes the symptoms of the "fear of floating" that characterizes monetary policy in Croatia and relates it to the consequences of dollarization. Section III reviews the literature on the exchange rate pass-through. Section IV deals with data issues. Section V presents the models used for the estimation and the results. Section VI concludes.

II. Fear of Floating and Dollarization

Most emerging market economies exhibit symptoms of what Calvo and Reinhart (2000) have termed "fear of floating." Due to the expected strong impact of the exchange rate on trade and inflation, central banks in small open economies can hardly be indifferent to its movements. In emerging market economies these concerns are typically compounded by large unhedged foreign currency exposures, which make the financial system vulnerable to exchange rate fluctuations. As a consequence, central banks in emerging markets feel generally uncomfortable with exchange rate movements and intervene frequently to smooth them. Even though there is no explicit peg and central banks claim their exchange rate is "floating", intervention can be substantial and, accordingly, the exchange rate relatively inflexible.

Croatia is a striking example of fear of floating. The CNB policy response to the wartime hyperinflation in the early nineties was to implicitly peg the currency to the deutsche mark. Even after stabilization was successfully achieved in 1994, monetary policy remained driven

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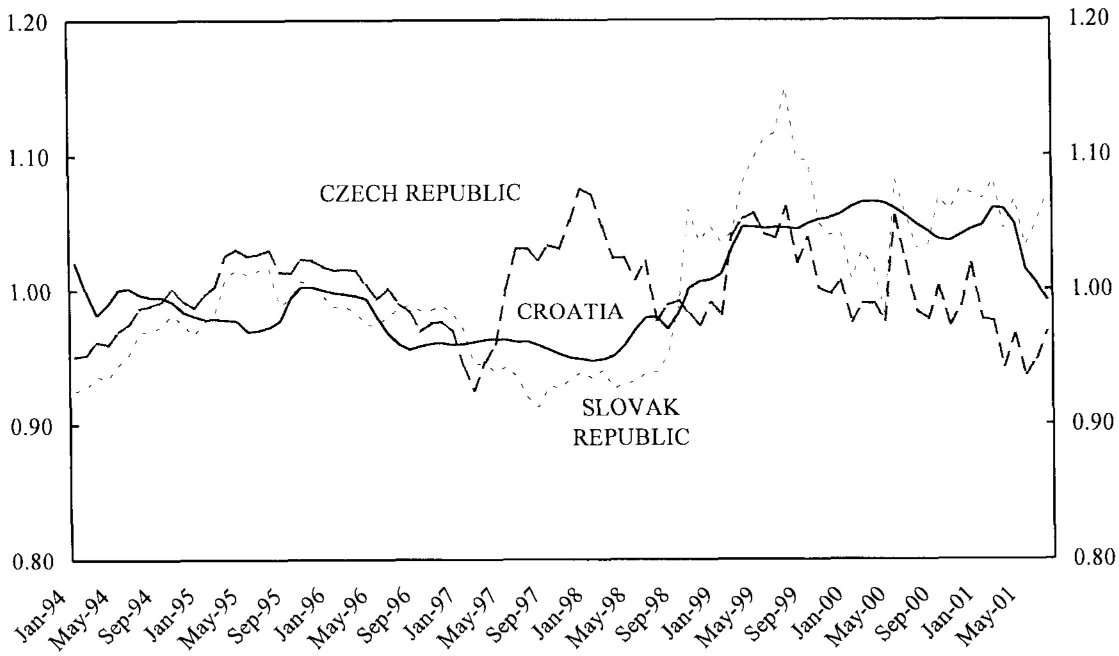
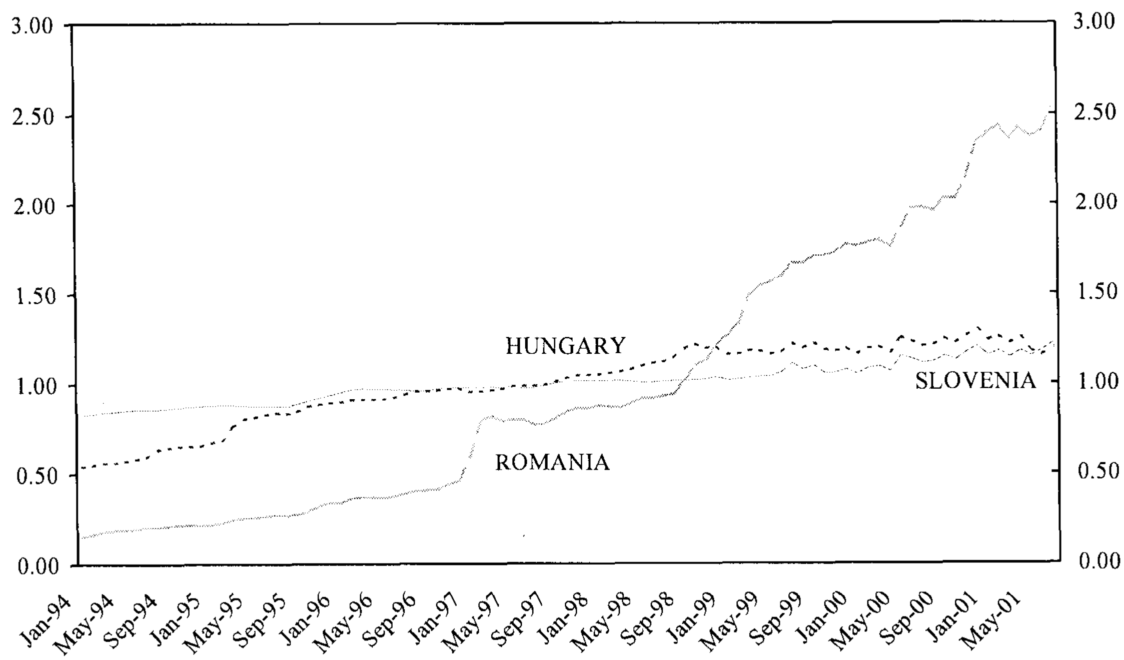
[5] The ratio of imports to GDP (average 1991-2000) is 0.53 in Croatia, 0.58 in the Czech Republic, 0.44 in Hungary, 0.32 in Romania, 0.65 in the Slovak Republic, and 0.60 in Slovenia.

Figure 1. DM Exchange Rate in Selected Countries, 1994-2001
(monthly average)





Source: CNB, IMF

Note: Data are divided by the mean.

by a strong concern for the exchange rate. [6] The tolerance of the CNB for exchange rate movements has been relatively low and intervention quite systematic. In this regard, it is interesting to compare Croatia with other transition economies with similar exchange rate arrangements, like the Czech Republic, Slovakia, Slovenia, Romania—all, like Croatia, "managed floats" in the IMF classification scheme—and a soft "peg" like Hungary. While the latter three have a history of continuous depreciation since 1994 (largely managed in the case of Slovenia and Hungary), the Croatian currency, like those of the Czech Republic and Slovakia, has been fairly stable over time and even within this group its movements have been remarkably smooth (Figure 1). This visual impression is confirmed by the volatility measures presented in Table 1, where Croatia displays by far the most inflexible exchange rate.[7] In the period January 1994-July 2001, the probability of the monthly percent change in the Kuna/DM exchange rate exceeding a 2.5 percent band has been a tiny 1.1 percent, much lower than for any other currency in the sample.

Table 1. Volatility of Selected Indicators in Some "Managed Floating" Exchange Rate Regimes, 1994-2001

Probability that the monthly percent change is in a +/- 2.5 percent band:					
	Exchange Rate 1/	Real Effective Exchange Rate 1/	Reserves 2/	Monetary Base 3/	Inflation 4/
Croatia	98.9	94.4	43.8	31.8	26.1
Czech Republic	81.1	90.0	58.4	59.1	36.4
Hungary	75.6	90.0	31.5	29.4	22.7
Romania	46.7	70.0	30.3	23.9	9.1
Slovak Republic	83.3	92.2	47.2	36.4	38.6
Slovenia	84.4	n.a.	49.4	28.4	22.7

Source: CNB, IMF

1/ Monthly averages, January 1994 - July 2001.

2/ End of period, January 1994 - June 2001.

3/ End of period, January 1994 - May 2001; Hungary: January 2000-May 2001.

4/ Annualized monthly rate of change in consumer prices, January 1994 - June 2001.

[6] For a review of the Croatian stabilization program see Sonje and Skreb (1997).

[7] This statistic has the important advantage over the variance of being robust to outliers, which are common in exchange rate movements (even in this sample: see Figure 1). See Calvo and Reinhart (2001).

When the exchange rate is fixed, shocks to money demand and expectations are accommodated by the central bank through purchases or sales of foreign exchange reserves. Thus, for a given distribution of shocks, the volatility of reserves and the monetary base are inversely related to the flexibility of the exchange rate. Table 1 shows that reserves and the monetary base in Croatia have indeed been the most volatile within the group of "stable" currencies, suggesting that the relative inflexibility of the kuna/DM exchange rate may in fact be the outcome of a deliberate policy by the CNB.[8] The low tolerance of exchange rate fluctuations, and the consequent volatility of the monetary base, is likely to be reflected in inflation. In fact, inflation volatility appears high relative to Czech Republic and Slovakia, even though a stable exchange rate and tight monetary policy have delivered low inflation in Croatia.[9]

What motivates this fear of floating? Most important are concerns about the consequences of dollarization, which is widespread in Croatia.[10] Foreign currency deposits represent more than 60 percent of broad money, a much larger percentage than for the rest of the sample (Figure 2). The origins of this phenomenon go back to the war of the early nineties to which both Croatia and neighboring Slovenia were subjected. In fact, the extent of dollarization in both countries was relatively large in 1994, when foreign currency was also used as a means of payment. While hostilities were soon over in Slovenia and disruption was relatively minor, the Croatian war continued till 1995 and was accompanied by extensive economic instability and hyperinflation. Even after macroeconomic stabilization was achieved in 1994, dollarization continued spreading until 1998, and has not significantly declined since.[11]

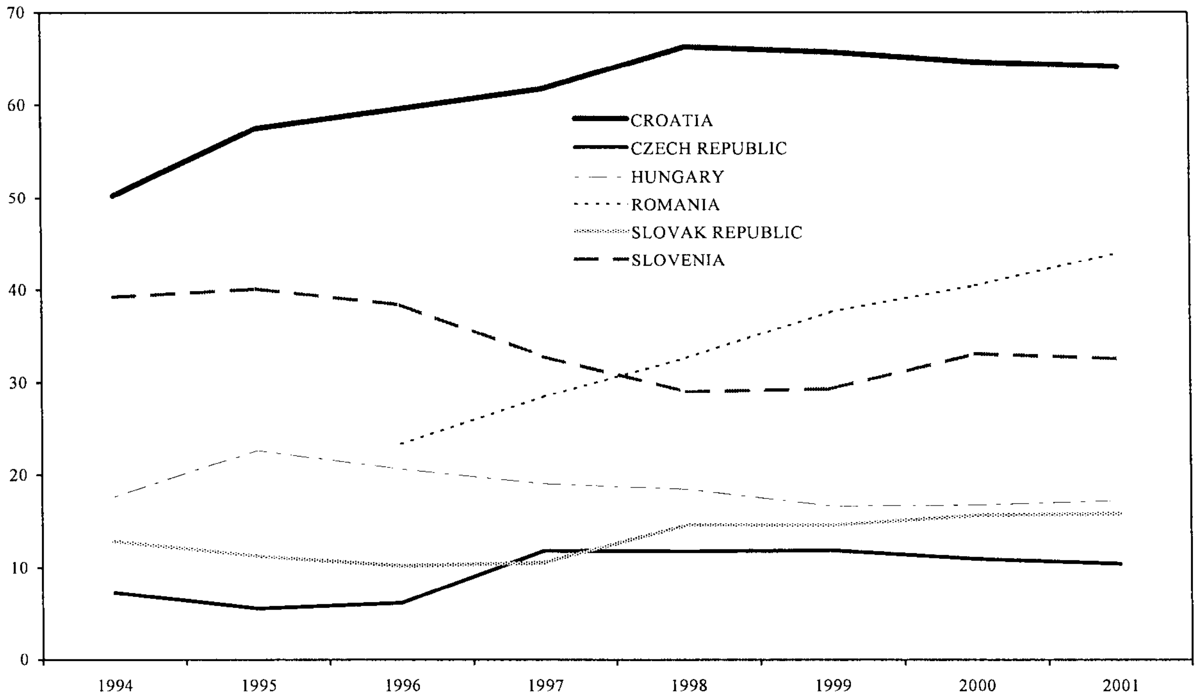
[8] Of course, there is also the possibility that Croatia had been subject to fewer and smaller exogenous shocks.

[9] Inflation was on average 5.1 percent in Croatia, 7.3 percent in the Czeck Republic, 9.1 percent in Slovakia, 10.4 percent in Slovenia, 16.9 percent in Hungary, and 87.2 percent in Romania.

[10] Or, more specifically, "euroization" due to the traditional link with the deutsche mark.

[11] This persistence, or "dollarization hysteresis", is certainly the result of long-lasting memory of the war-time hyperinflation. However, the very policy followed by the CNB, by excessively restricting the volatility of the nominal (and real: see Table 1) exchange rate, may have contributed to increase the attractiveness of foreign currency. Using a CAPM model, Ize and Levy Yeyati (1998) show that, for a given variance of inflation, lower exchange rate volatility increases the hedging benefits of foreign currency assets.

Figure 2. Dollarization in Selected Countries, 1994-2001
(foreign currency deposits as a percentage of broad money)



Source: IMF

Note: Data are end-December for 1994-2000, end-May for 2001.

Under these circumstances, any independent monetary policy is faced with serious risks of financial instability and its effectiveness in controlling inflation is likely to be severely limited. Sudden movements in the exchange rate may damage the balance sheet of firms and individuals, causing an increase in non-performing loans and thus endangering the stability of the banking system. Moreover, the impact of monetary tightening may be weakened if the consequent exchange rate appreciation improves the financial position of residents with a large proportion of foreign currency debt. In the extreme case of "real dollarization", when prices and wages are set in foreign currency, monetary policy becomes completely ineffective and the only possible strategy is to target the exchange rate (Ize and Levy Yeyati, 2001).

Nowadays dollarization in Croatia is mainly motivated by asset substitution, as residents maintain a large proportion of their savings in foreign currency and banks provide loans that are either foreign currency-denominated or indexed to foreign currency. Casual observation indicates that many prices, mainly of property and consumer durables, are to some extent indexed to the exchange rate. This would argue for a large pass-through coefficient.

III. The Pass-Through Literature

The relationship between the exchange rate and the price level has received a large amount of attention since the breakdown of Bretton Woods. Recent currency crises have only added. Very few contributions have analyzed the complete pass-through, i.e., the effect of an exchange rate change on various domestic price measures along the production chain (including export/import prices and a measure of consumer inflation), however. Instead, most research focuses on particular segments, such as import/export price pass-through, or synchronization of different price level measures.

The direct impact of exchange rate movements occurs through prices of internationally traded goods. Goldberg and Knetter (1997) survey the literature on pass-through to import prices. An important question that arises in this context is why the pass-through might be incomplete. Many theoretical explanations have been found to this question, see Menon (1995), who surveys a large number of contributions. Most prominently, Dornbusch (1987) and Krugman (1987) show that a less than one-to-one transmission can be explained by imperfect competition, or "pricing-to-market". Foreign producers adjust their mark-up to maintain a stable market share in the domestic economy. This strategic behavior can in principle drive the rate of pass-through to zero. Gosh and Wolf (2001) show that imperfect short-run pass-through may also arise from menu costs. Few papers look at a larger sample of countries. Exceptions are Borensztein and De Gregorio (1999) who examine currency crises, and Goldfajn and Werlang (2000) who show in a panel framework that the main determinants of the extent of pass-through are the cyclical state of the economy, the initial over/undervaluation, the initial rate of inflation, and the degree of openness of the economy. They also find that pass-through peaks after twelve months.

McCarthy (2000) analyzes in a stationary VAR framework (incorporating a recursive distribution chain of pricing) the impact of exchange rate changes and import prices on producer and consumer prices. In his sample of nine developed countries, he finds that the exchange rate has a rather limited and insignificant effect on consumer prices while import prices have a stronger effect. Clark (1999) examines the responses of prices at different stages of production in different context, namely in response to a domestic monetary policy shock. Ross (1998) provides an assessment of Slovenia, using a recursive model of the economy. Kuijs (2001) analyses the transmission of monetary policy in Slovakia using a structural cointegrated VAR. The causality issue is addressed in a number of papers that deal specifically with the Australian economy:[12] A significant share in the world market for a specific commodity might imply potential feedback from the domestic price level into the world price level, therefore questioning the small open economy assumption. This is found to be true with regard to Australia for a number of commodities.

[12] See among others Webber (1997) and Dwyer et al (1993).

From a technical point of view, empirical analysis has to deal with the fact that most time series to be analyzed are certainly non-stationary. This invalidates OLS estimation, creating the problem of "spurious regressions". By employing the variables in first differences this trap can most probably be avoided,[13] but the "level information" is lost. Few papers so far use cointegration techniques to account for non-stationarity and causality at the same time: Kim (1998) presents US evidence in favor of a causal relationship running from the exchange rate to prices in a cointegrated VAR framework. He finds a long-run pass-through coefficient of 0.24, but does not address short-run adjustment issues. Murgasova (1996) applies the Johansen Maximum Likelihood (ML) method to analyze the effects of the Spanish peseta's devaluation during the ERM crisis in 1992–93. She finds a one-to-one pass-through to import prices but only a 10 percent pass-through to the CPI. Low pass-through is explained by the offsetting cyclical position of the economy. Dellmo (1996) focuses on the relationship between Swedish price level measures in an I(1) framework, taking into account factors that might limit expected similarities, such as varying profit margins and productivity. Juselius (1999) analyzes price level convergence in a full-fledged I(2) framework.

In this paper, we will initially follow the former road. A standard VAR (in first differences) will be estimated. The underlying model of the distribution chain translates into a recursive structure of the variance-covariance matrix. This in turn will enable us to identify shocks stemming from exchange rate changes and their effects on Croatian inflation. The structure is a stripped-down version of McCarthy (2000), who employs a production chain model inspired by Blanchard (1983) and Christiano et al. (1997). The reduction in complexity (number of variables) is mainly due to lack of data for Croatia and will be discussed in more detail below. A major drawback of this set up is the imposed causality from the exchange rate to prices. Assuming that prices are sticky in the short run, for purchasing power parity to hold, the exchange rate is supposed to move. This implies that causality could also run the other way. Given that the analysis is based on monthly observations, we feel confident however about imposing the condition there is no-contemporaneous-feedback. Another weakness of this approach is the standard criticism to estimating differenced VAR systems mentioned above, namely that the level information is lost, and that therefore results (here impulse responses) will eventually lack statistical significance.

In order to assess in more detail the above-mentioned limitations, the long-run relationship between the exchange rate behavior and the price level will be investigated using a cointegration approach. Results from this exercise have to be handled with caution, however, since the observation period is rather limited (roughly 6 years), and monthly observations are subject to a high "noise-to-signal" ratio.

[13] This does not hold if level variables are in fact integrated of order 2, or I(2), since in that case, the first difference is still non-stationary.

IV. Data Issues

In this section, we will discuss data-related issues. The monthly time series are mostly taken from the IFS database of the IMF or provided by the CNB.[14] Although observations generally start in January 1992, the sample will be restricted to start in January 1994, given that stabilization was only achieved by end-1993. The observations for M4 start in June 1994. The estimation period is limited by the availability of the raw materials price index, which ends on January 2001. In this section, we first highlight the stylized facts related to exchange rate pass-through in Croatia. After that, we turn to stationarity issues. Most of the underlying time series employed seem to be non-stationary (in levels). Finally, issues of causality will be discussed.

A. Data Description

In Figures 3 and A1 we present preliminary evidence of our recursiveness assumption. Figure 1 displays the time series (in log-levels) for the kuna-over-deutsche mark average monthly exchange rate (HRK/DEM) and the nominal effective exchange rate (NEER, as calculated by the IMF), together with the two price series, the manufacturing price (MPI) and the retail price index (RPI), all adjusted for mean and range.[15] Visual inspection indicates that the RPI is hardly responding to moves in the exchange rate, in fact, the series seems almost to be trend stationary. The manufacturing price index, which reflects industry and therefore "upstream" behavior, seems to follow the exchange rate, albeit very slowly. The significant rise in the exchange rate during the second half of 1995 seems to be reflected in manufacturing prices roughly one year later. The sustained depreciation of the Kuna starting in early 1998 instead seems to lead to a rise in the price level, again roughly 12 to 15 months later. This preliminary assessment is confirmed by Figure A1 in the Annex, where cross-correlations between the change in the exchange rate and subsequent changes in the price indices are presented.

Figure A1 indicates that the correlation between the change in the exchange rate and the change in the MPI (lower panel) is positive from the ninth until the 26th (lagged) month, with a peak between months 13 and 16, i.e., after roughly one year. The peak correlation coefficient is not extremely high (0.22), however, and only marginally significant.[16] This might be due to the fact

[14] For a description of data sources, see section C in the appendix.

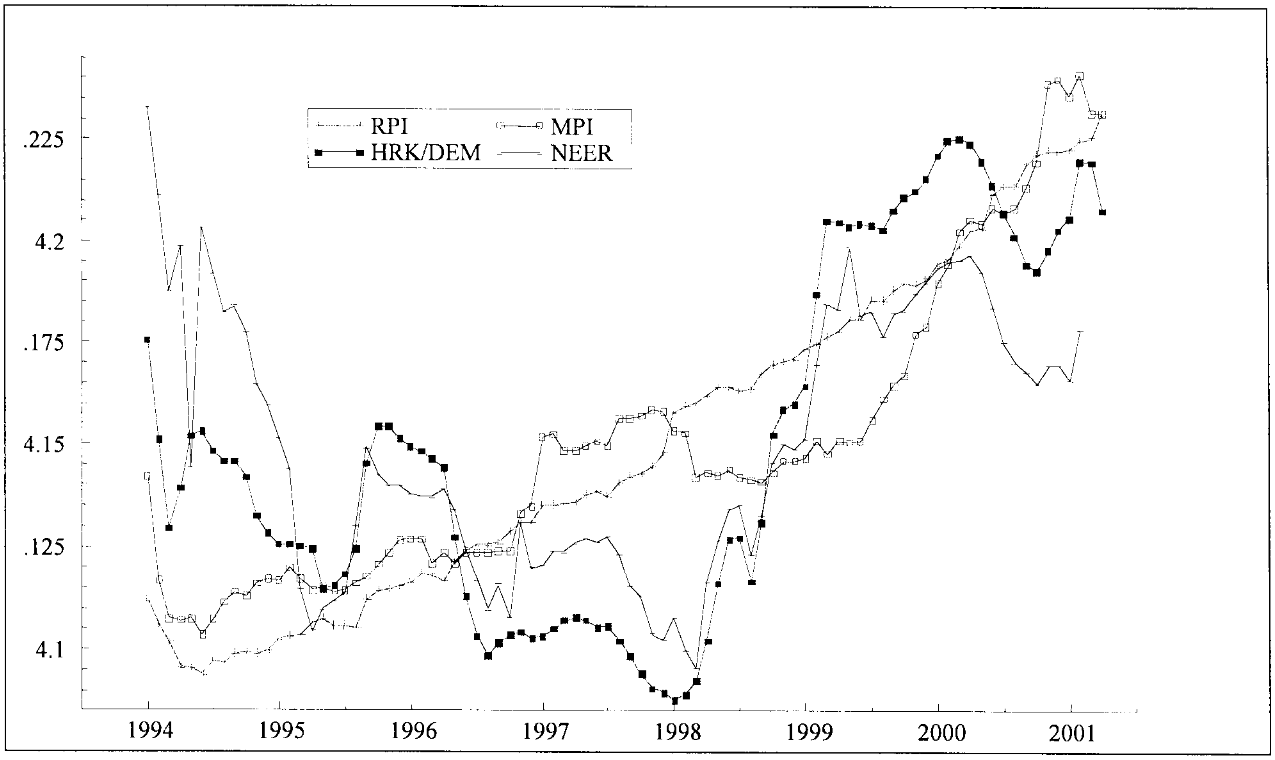
[15] Note that the NEER and the HRK/DEM exchange rates behave in a rather similar fashion. This might be due to the fact that the major trading partners of Croatia are mostly Euro/former ERM countries (Germany and Italy accounting for roughly 20 percent of exports and imports each). the major trading partners of Croatia are mostly Euro/former ERM countries (Germany and Italy accounting for roughly 20 percent of exports and imports each), the NEER and the HRK/DEM behave rather similarly. Furthermore, "dollarization" in Croatia is DEM-based; therefore we focus in what follows on the exchange rate against the deutsche mark.

[16] The horizontal lines indicate two standard deviations.

that the MPI contains already less items subject to exchange rate changes than other, more upstream price indices. In Croatia, no sudden devaluation has taken place, and therefore, the attribution of price changes to particular exchange rate movements is more difficult. The cross-correlation exercise for the RPI confirms the visual impression conveyed by Figure 3: Retail prices are hardly affected by exchange rate movements, the correlation often changes sign, and is hardly higher than 0.10 (in absolute terms).

Figure 3. Exchange Rates and Price Indices

(Logs of levels)



B. Granger Causality

In this section, pairwise Granger causality (GC) tests based on the single time series "by lag" (standard F-tests) are provided. Section V.A. presents Wald-type GC tests in the estimated VAR, taking into account all lags of the respective variable. The first type of causality analysis (see Table B2 in the appendix) reveals four empirical facts: Firstly, the null of no granger causality running from the change in broad money (DM4) to the output gap is firmly rejected at almost all lags. This can be interpreted as a traditional demand side argument, where strong money growth increases real balances and causes demand to exceed supply. Secondly, changes in the retail price index (DRPI) helps to explain the output gap. Thirdly, and perhaps even more surprisingly, the hypothesis that the output gap does not Granger cause the raw materials price index computed by the HWWA is strongly rejected. We do not offer any explanation for this statistical phenomenon and interpret it as spurious. And finally, the same raw materials index seems to "lead" the manufacturing price index. This result fits the distribution chain model chosen as rationale for the recursive structure of the empirical model: the proposed ordering {DHWWA→HPOGAP→DKDAV→DMPI→DRPI→DM4} does not encounter strong empirical objections. At one lag, only the hypothesis that DKDAV leads DM4 is rejected at the 5 percent-level. Experimenting with a different ordering (grouping DM4 before or after HPOGAP) did not change the results substantially. Note that the change in the exchange rate is "leading" MPI inflation only significantly at 3 lags and the RPI not at all.

C. Stationarity

In this section, the order of integration of the data is investigated. In order to estimate the unrestricted VAR,[17] the time series employed are assumed to be stationary. Table B2 (in the appendix) presents the characteristics. Two standard types of unit root tests have been applied, augmented Dickey-Fuller (ADF), and Phillips-Perron (PP).[18] These tests were carried out for the period starting in January 1994 (with the exception of the monetary aggregate M4, which was available from 1994:6 only) until 2001:05 (where available). For this type of test, the null hypothesis of a unit root (UR) is tested against the alternative of stationarity.[19] It can be seen that in principle all variables of interest are non-stationary in levels, but stationary in first differences. The two exceptions are the RPI and the industrial production (IP)/the output gap. It comes as a slight surprise that a price level time series can be considered stationary, given that most empirical evidence from other countries points to higher orders of integratedness, including even potential non-stationarity of the inflation rate, i.e., the first difference.[20] The initial visual impression of trend stationarity is confirmed by the fact that, if the UR tests do not include a time trend, stationarity is rejected. The series for industrial production by itself does not exhibit strong trending behavior, neither does the output gap; see Figure A2 in the Appendix.

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[17] i.e., without long-run cointegration restrictions.

[18] The ADF test based on Dickey and Fuller (1979) augments the standard DF test by including additional lag terms to account for higher order residual autocorrelation. The PP test, based on Phillips and Perron (1988), applies a non-parametric correction to the t-statistic of the coefficient in the estimated AR(1) process.

[19] This implies that in the potential presence of I(2) trends, testing should start with first differences, and proceed then to test the variables in levels to guarantee a statistically correct testing sequence, Banerjee et al. (1993), pp.119s. Given that visual inspection did not indicate the presence of I(2) trends, we refrain from this practice.

[20] See Juselius (1999), who analyzes price indices integrated of order 2, or I(2).

In the first part of the empirical investigation, the recursive VAR, the series will be included in first differences with the exception of the output gap. The cointegrated VAR employs only three time series (the exchange rate and the two price indices, all in levels).

V. Empirical Analysis

In this section, we will present the models and report the results from the stationary VAR modeling approach and the non-stationary, cointegrated VAR.

A. A Recursive Approach

Set-up

Following McCarthy (2000), we assume a recursive ordering, under which international supply side shocks (approximated by the HWWA commodity price index) and demand shocks (as given by a measure of the output gap[21]) are exogenous to the exchange rate shock in period t. They are determined in each period by the expectations of the previous period and an error,

$$\pi_t^{wp} = \underset{t-1}{E}\left(\pi_t^{wp}\right) + \varepsilon_t^{wp}$$

$$y_t = \underset{t-1}{E}\left(y_t\right) + a_1\varepsilon_t^{wp} + \varepsilon_t^{y}$$

$$\Delta e_t = \underset{t-1}{E}\left(\Delta e_t\right) + b_1\varepsilon_t^{wp} + b_2\varepsilon_t^{y} + \varepsilon_t^{\Delta e}$$

where π_t^{wp} is the inflation in world commodity prices, y_t measures the output gap, Δe_t the change in the exchange rate, and $\varepsilon_t^{\cdot}$ are the respective shocks which occur at each stage. This exchange rate shock then feeds into domestic inflation, first at the manufacturers' level, and then at the retail level.

$$\pi_t^{mpi} = \underset{t-1}{E}\left(\pi_t^{mpi}\right) + c_1\varepsilon_t^{wp} + c_2\varepsilon_t^{y} + c_3\varepsilon_t^{\Delta e} + \varepsilon_t^{mpi}$$

$$\pi_t^{rpi} = \underset{t-1}{E}\left(\pi_t^{rpi}\right) + d_1\varepsilon_t^{wp} + d_2\varepsilon_t^{y} + d_3\varepsilon_t^{\Delta e} + d_4\varepsilon_t^{mpi} + \varepsilon_t^{rpi}$$

The main differences with respect to the above-mentioned paper are that we do not include a measure of import prices due to lack of data, and that we do not explicitly model the behavior of the central bank. Given that Croatia is a small open economy without significant power on the world market, we expect the pass-through on import prices to be complete over a rather short time horizon. McCarthy adds two more variables, namely the interest rate (which is supposed to respond to the aforementioned variables in a way similar to a central bank "rule") and money stock growth, to reflect money demand behavior. In this paper, the former variable is not included, since there is no well-functioning money market in Croatia, i.e., the interest rate would not reflect market-type behavior. The latter variable is included in a non-standard way,

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[21] See Babic and Stucka (2000) for an overview of measures of the Croatian output gap.

because of the peculiar structure of the Croatian money supply, owing to the high DEM-dollarization. The money stock defined as M4 therefore reflects the central bank's behavior (M0/M1) as well as private sector decisions, both of which are assumed to adapt to the above-mentioned variables:

$$\Delta m_t = \underset{t-1}{E}(\Delta m_t) + e_1\varepsilon_t^{wp} + e_2\varepsilon_t^{y} + e_3\varepsilon_t^{\Delta e} + e_4\varepsilon_t^{mpi} + e_5\varepsilon_t^{rpi} + \varepsilon_t^{\Delta m}.$$

Starting from a structural form represented by a set of linear dynamic equations of the form

$$A_0 X_t = A(L) X_{t-1} + B\varepsilon_t,$$

where X_t is the p(=6)-dimensional vector containing the variables of interest, A_0 describes the contemporaneous relations between the variables, A(L) is a finite-order matrix polynomial in the lag operator L, and ε_t is a vector of (interpretable) structural disturbances, taken from the equations describing the system (see above) with covariance matrix Σ_ε. Non-zero off-diagonal elements of B would allow shocks to affect more than one variable. As it is well known, the structural model is not observable. Under mild conditions (A_0 invertible), we can express the p-dimensional stationary autoregressive process X_t in the following (reduced-form) way:

$$X_t = A_0^{-1} A(L) X_{t-1} + e_t,$$

where the VAR residual vector $e_t = A_0^{-1} B\varepsilon_t$ is n.i.i.d. with full variance-covariance (VCV) matrix Σ_e. From this, we can derive the relation between the VCV matrices of (unobserved) ε_t and (observed) e_t :

$$E(e_t e_t') = A_0^{-1} B E(\varepsilon_t \varepsilon_t') B' A^{-1}.$$

In the sample, it holds that $\hat{\Sigma}_e = \hat{A}_0^{-1}\hat{B}I\hat{B}'\hat{A}_0^{-1}$. Identification requires that restrictions be imposed on A, B. The Cholesky decomposition, originally proposed by Sims (1980), is the most well known way (and the one followed by McCarthy (2000)). Under this strategy, the matrix A is assumed to be lower-triangular, whereas B is assumed to be diagonal:

$$A = \begin{pmatrix} 1 & 0 & \cdots & 0 \\ a_{21} & 1 & & \\ \vdots & & \ddots & \\ a_{n1} & a_{n2} & \cdots & 1 \end{pmatrix}, \quad B = \begin{pmatrix} b_{11} & 0 & \cdots & 0 \\ 0 & b_{22} & & 0 \\ \vdots & & \ddots & 0 \\ 0 & 0 & 0 & b_{nn} \end{pmatrix}.$$

In this scheme, just-identification of shocks depends on the variable ordering. The imposed recursiveness has been the main criticism brought forward in the literature. Indeed, a (block-) recursive structure implies that the "degree of endogeneity" rises along the variable ordering. In our case, this argument against Cholesky can be flipped around: production chain and limited participation models precisely rationalize this type of structure.

Once the recursive model has been estimated, a number of exercises can be accomplished. Variance decompositions show for each variable the ratio of the forecast error variance that is attributable to its own shocks and to shocks stemming from other (upstream) variables. Impulse response functions show the estimated response of each variable to an impulse in one of the innovations. The impulse responses of MPI and RPI to the exchange rate will provide estimates of he importance of the exchange rate channel for domestic inflation at different stages. As it was mentioned above, VAR modeling implies that contemporaneous correlations are reflected in the cross-equation residual correlation. The Cholesky factorization sets to zero the residual correlation between a given variable and another variable prior in the (causal) ordering.

Results

The VAR in first differences is estimated with three lags, accounting for the additional noise present in monthly series compared to quarterly observations. The 6*3=18 calculated roots of the characteristic polynomial are located all within the unit circle, hence the system is stable.[22] According to visual inspection, the residuals of almost all series display a number of significant statistical outliers,[23] such that we do expect significant non-normality.[24] To check whether this description of the data is consistent with the assumption of white noise errors, multivariate serial correlations of the residuals are calculated. Using a Lagrange Multiplier test (distributed as $\chi^2(36)$), the null hypothesis of no autocorrelation could not be rejected at the 10 percent level for any lag <10 with the exception of lag 6 (significant at the 10, but not at the 5 percent level). In Table 2, univariate diagnostic statistics are presented. Due to the strong non-normality of the price indices, multivariate normality is rejected at the 5 percent-level, as well. Table B3 presents the matrix of residual correlation. Off-diagonal elements are rather close to zero, such that no contemporaneous correlation is being ignored by the VAR.

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[22] The biggest root is a complex pair at 0.77.

[23] In Figure A3, the doted line indicates two standard deviations.

[24] Note that adding lags to the VAR could not remove the substantial outliers.

Table 2. DiagnosticTests.

Time series	Skewness	(prob.)	Kurtosis	(prob.)	JB	(prob.)
DHWWA	-0.064	(0.82)	**1.634**	**(0.01)**	5.965	(0.05)
HPOGAP	-0.387	(0.17)	2.046	(0.09)	4.781	(0.09)
DKDAV	0.053	(0.85)	3.811	(0.15)	2.120	(0.35)
DMPI	**0.929**	**(0.00)**	**4.393**	**(0.02)**	**17.08**	**(0.00)**
DRPI	**0.868**	**(0.01)**	3.631	(0.26)	**10.81**	**(0.01)**
DM4	0.357	(0.20)	2.185	(0.15)	3.723	(0.16)

Note: Series are log first differences with the exception of the output gap. Skewness and kurtosis of a normal distribution are 0 and 3, respectively. Under the null of normality the Jarque-Bera test for normality is distributed as χ^2 with 2 degrees of freedom.

A second battery of Granger-causality tests, carried out in the estimated VAR framework (see Table B4 in the appendix), confirms in principle the preliminary assessment given above. Changes in broad money seem to have a causal effect on the output gap, the raw material price influences the MPI, but not the RPI, and exclusion of the kuna-deutsche mark exchange rate as a "cause" for the price levels cannot be rejected at conventional levels.

We now turn to variance decompositions and impulse response functions for the VAR estimated. While the former breaks the variation in an endogenous variable down to the component shocks to the endogenous variables in the VAR, the impulse response functions trace the effects of a shock to one endogenous variable on the other variables through the dynamic structure of the VAR. The variance decompositions over ten periods for changes in the exchange rate and the price indices are presented in Figures A4-A6. Generally speaking, variances of all three variables are mostly explained by their own innovations at all horizons. Comparing the two price indices, it is noted that the changes in MPI can to some extent be attributed to the change in the raw materials index and to the exchange rate. Interestingly, the latter has a particular contemporaneous effect on the MPI variance. In the case of the downstream RPI, the importance given to other innovations is negligible. This confirms the impression that the intermediate price index is to some extent influenced by changes in the exchange rate while the more consumer-oriented retail price index does not respond to external factors.

The impulse response functions are reported in Figures A7 and A8. The first figure reports (by column) the responses to a one standard deviation innovation to the change in the exchange rate, the manufacturing inflation and retail inflation. Responses are hardly significant for any pair of variables. This lack of significance is attributed to three main factors: firstly, the fact that observations span only seven years. Secondly, observations are monthly and therefore subject to a high signal-to-noise ratio. Thirdly, low variation in the data, particularly the exchange rate, reduces estimation accuracy. In Figure A7, the MPI shows a positive but insignificant reaction to an exchange rate shock. The RPI instead does hardly react at all to changes in the exchange rate. The cumulative (i.e., level) response to a disturbance in the first difference (Figure A8) offers slightly more insight: Own effects are significant over the whole period for all variables, and a clear (but almost insignificant) level effect can be seen in the MPI, not in the RPI, though. Again, the impression is confirmed that the exchange rate has a lasting effect on the manufacturing price. The short run evolution of RPI instead seems to be driven by variables not modeled in the present setting.

Although these results do not allow strong statements about short-run pass-through effects due to the lack of statistical significance, they are roughly comparable to what has been found in a similar set-up for a number of freely floating developed countries by McCarthy (2000). There, import prices do have a significant impact on more downstream measures of the price level, the exchange rate however does not. This result cannot be verified in our setting, given that such an index is not published for Croatia. The lack of significant pass-through at the second stage can be interpreted as indirect evidence of the limits discussed above: pricing to market, but also institutional constraints, such as administered prices, reduce the effect of an exchange rate movement. The limited information contained in the above results suggests a different approach, taking into account the level information present in the data. In the following, we will focus on the long-run relationship between the exchange rate and the final price index, in our case retail prices.

B. A Cointegration Approach

The cointegrated VAR approach can deliver valuable additional information for a number of reasons. For one, the cointegration property is invariant to an augmentation of the information set, i.e., if non-stationary time series appear cointegrated in a small model, they will still be so in a bigger model.[25] This allows the estimation of a model consisting only of exchange rate and price indices in levels, resulting in more precise coefficient estimates. Second, the richer theoretical structure allows to impose restrictions on long and short run and to analyze both types of dynamics. Thirdly, the issue of Granger causality is analyzed in a more direct framework. Results of this causality test are expected to differ from the above results since only the three time series of interest are modeled.

Multivariate misspecification tests indicate that when estimated with three lags as above, in the three-variable cointegrated VAR system autocorrelation at lag 1 is significant at the 5 percent-level, but insignificant at lag 4. The smallest system free from residual autocorrelation consisted of four lags and allows the preliminary assessment of the single time series as reported in Table 3.

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[25] The common trends analysis however might be substantially different.

Table 3. Preliminary Tests

			Time series			
R	dgf	$\chi^2(95\%)$	KDAV	MPI	RPI	trend
Test for exclusion ($LR \sim \chi^2(r)$)						
1	1	3.84	**12.05**	**10.24**	**11.93**	**9.87**
2	2	5.99	**14.26**	**15.58**	**19.00**	**16.36**
Test for stationarity ($LR \sim \chi^2(p-r)$)						
1	3	7.81	**18.58**	**20.03**	**17.35**	
2	2	5.99	**8.79**	**9.25**	5.48	
Test for weak exogeneity ($LR \sim \chi^2(r)$)						
1	1	3.84	**6.81**	0.04	**5.85**	
2	2	5.99	**15.27**	2.73	**14.12**	

Note: r is the cointegrating rank, p the number of variables (3), and dgf the degrees of freedom. Entries indicate the χ^2 test statistic, a value higher than the 95 percent critical value (χ^2(dgf)), indicates rejection of the null. Rejections are printed in bold.

The time series are all fundamental to the system estimated and stationarity is rejected for all of them. Tests on the individual series indicate that in the smaller model, weak exogeneity of the intermediate price index MPI cannot be rejected for all choices of r, the cointegration rank. Multivariate tests indicate no autocorrelation at conventional levels for lag 1 and 4 (p-values of 0.96 and 0.40, respectively), normality for the VAR residuals is, again, strongly rejected.[26] Univariate ARCH test statistics are not significant. Hence, the model seems well specified. Therefore, the cointegrated VAR is estimated imposing exogeneity of the MPI. Note that imposing exogeneity of the MPI lowered the autocorrelation test statistics slightly (to 0.86 and 0.18, respectively). The contemporaneous correlation between the two endogenous series amounts to -0.032. Examination of an even smaller system consisting of only the exchange rate and the MPI resulted in no cointegration between the two series.

Set Up

The cointegrated VAR(k) framework can be represented in error correction form (ECM) as follows:

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[26] Note that the validity of results in the cointegration framework rests on the assumption that the residuals be i.i.d, not n.i.i.d. It is important therefore to account for residual autocorrelation, but not necessarily for normality.

$$\Delta X_t = \Pi X_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \Phi D_t + \mu_0 + \mu_1 t + \varepsilon_t,$$

where X_t is a p-dimensional autoregressive process, k is the lag length, ε_t is an i.i.d. error with mean zero and variance Ω, $\Pi = \sum_{i=1}^{k} \Pi_i - I$, $\Gamma_i = -\sum_{j=i+1}^{k} \Pi_i$, and D_t contains seasonal and intervention dummies. Under the I(1) hypothesis that rank(Π)=r<p, we can decompose $\Pi=\alpha\beta'$ where α,β are p×r matrices of rank r, and $\alpha'_\perp \Gamma \beta_\perp$ has full rank (p-r), where $\alpha_\perp, \beta_\perp$ are the orthogonal complements of α, β and where $\Gamma = \sum_{i=1}^{k-1} \Gamma_i$. The trend is restricted to the cointegrating space, i.e., $\alpha'_\perp \mu_1 = 0$, since we do not observe quadratic trends in the data. The moving average representation of this I(1) process defines the data generating process for X_t as a function of the errors ε_t, the initial values A_0, and the variables in D_t. It is given by:

$$X_t = C\left(\sum_{i=1}^{t} (\varepsilon_i + \Phi D_i) + \mu_0 t\right) + C^*(L)(\varepsilon_t + \mu_0 + \mu_1 t + \Phi D_t) + A_0,$$

where the impact matrix $C = \beta_\perp (\alpha'_\perp \Gamma \beta_\perp)^{-1} \alpha'_\perp$. $C^*(L)$ is a finite polynomial in the lag operator L, A_0 is a function of the initial values.

The cointegrating vectors are estimated by reduced rank regression of ΔX_t on (X_{t-1},t), corrected for lagged differences and the constant, see Johansen (1996), Theorem 6.2. The estimated model disposes of 84 observations and contains a trend restricted to the cointegrating space as well as constant and seasonal dummies, leaving 58 degrees of freedom.

Results

The trace test statistic[27] for the cointegrating rank reported in Table 4 is estimated using Johansen's full information maximum likelihood procedure.

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[27] The trace statistic seeks to identify how many eigenvalues are significantly different from zero, in other words the number of dimensions of the cointegrating space. H_0 indicates the null hypothesis, the alternative being r=i+1. Here, λ_i gives the eigenvalues obtained in the process of the maximization of the likelihood function. They correspond to squared canonical correlation coefficients and indicate, roughly speaking, the degree of correlation between the stationary part of the system and the potentially stationary cointegrating vector. λ_1=0.271 indicates therefore a correlation of approx. 55 percent. Trace indicates the test statistic from the CATS output (rejections in bold), and trace95 gives the critical value, more precisely the 95-percent quantile of the likelihood ratio test for the cointegrating rank, taken from Johansen (1996), Table 15.4.

Table 4. Test for the Cointegrating Rank

H_0	λ_i	Trace	Trace95
$r = 0$	0.271	**38.28**	25.47
$r < 1$	0.131	11.79	12.39

The hypothesis $r = 0$ is strongly rejected, while the second hypothesis is not. Further evidence for $r = 1$ can be obtained from the estimated adjustment coefficients α, which indicate significant adjustment (in the sense of error correction) only in the first cointegrating vector. Also the eigenvalues of the companion matrix point to $r = 1$.[28] In the remaining analysis, we will therefore assume one cointegrating vector. Table 5 gives the unrestricted estimates of the cointegrating relationship β and the adjustment coefficient α, normalized on the RPI.

Table 5. Unrestricted Estimates [29]

Variable	β	α
KDAV	-0.327	0.211
		(4.123)
RPI	1	-0.187
		(-3.603)
MPI (exog.)	-0.404	
Trend	-.001	

[28] In the companion matrix, eigenvalues that are "close" to unity correspond to the (p-r) remaining stochastic trends of the system. In the present case, we find the highest unrestricted eigenvalues at 0.96 and 0.75, while for r = 1 no remaining eigenvalue comes close to the unit circle, indicating that there is no further non-stationary trend.

[29] T-values for the adjustment coefficients are in parenthesis. For the β-coefficients estimates, no *t* values could be obtained, since the system is not identified in an econometric sense. The coefficients are insofar significant as that the omission of any single or combination of variables from the system resulted in a rejection at the 1 percent-level of the respective restricted β-vector.

For easier interpretation, we can rewrite the long-run equilibrium relation as:

$$RPI=0.327\ KDAV+0.404\ MPI\ +0.001\ t.$$

Hence, the retail price index moves positively with the exchange rate and the MPI over time, the latter having a bigger effect. The coefficient to KDAV could be interpreted as long-run pass-through coefficient, indicating that a 10 percent devaluation results in a 3.3 percent rise of the retail price level. This cannot be regarded as a "rule", since this result does not emerge as a "deep" parameter —in the Lucas sense — from a structural model. It indicates a significant pass-through, higher than in other countries with a lower degree of dollarization.[30] However, its size does not confirm the widespread indexation of wages and prices that is generally reported by casual observation. On the other hand, the manufacturing price index has a long-run coefficient of roughly 0.4, indicating that 40 percent of a change in the MPI feed into the RPI. The significant adjustment coefficient for the RPI has the correct sign and indicates significant error correction. Note however that also the exchange rate is significantly adjusting to disequilibria. This is consistent with the above-mentioned view that the exchange rate is (partly) endogenous to Croatian monetary policy, i.e., part of a managed (non-)float.

VI. Conclusions

This paper focuses on the exchange rate pass-through in Croatia, i.e., the extent to which changes in the exchange rate feed into domestic price level indices. Measures of pass-through are important for a number of reasons. Monetary authorities tend to see the exchange rate as one of the major channels of the monetary transmission mechanism. Especially for small open economies, such as Croatia, it is extremely relevant for the policy maker to assess to what extent domestic inflation is affected by the exchange rate. A large pass-through, if associated with diffused indexation of wages and prices ("real dollarization"), would represent a serious constraint on the effectiveness of monetary policy and would require an exclusive focus on the exchange rate. The conduct of monetary policy in Croatia, which can be characterized as "strict exchange rate targeting", seems to be based on this assumption.

The paper estimates the pass-through coefficient using two different methods. The first is a stationary, recursive vector autoregressive (VAR) system where exchange rate shocks feed into the manufacturing price index and the retail price index. Although the intermediate price index initially seems to respond significantly to exchange rate movements as well as to movements in commodity prices, the retail price index does not. Even though the lack of an import price index may somewhat affect the precision of estimates for Croatia, this evidence is consistent with what McCarthy (2000) finds for most advanced economies. The second method is a cointegrated VAR, where the information contained in the non-stationary data is fully exploited.

[30] The comparable coefficient found by Kuijs (2001) for Slovakia is 0.2.

Concentrating on the long run, a clearer result emerges. The pass-through from the exchange rate to the intermediate price index cannot be captured, but a pass-through coefficient of roughly 0.3 is obtained for the retail price index. This qualitative difference in results may be due to the fact that the recursive structure defines the pass-through rather narrowly as a causal consequence to disturbances in the exchange rate, while in the cointegration approach the pass-through relationship can be considered a pattern of the macroeconomic outcome, not necessarily yielding a causal interpretation. In this sense, the recursive approach is more demanding. In any case, the estimated coefficient can hardly be regarded as proof of diffuse exchange rate indexation of prices and wages.

There are reasons to be cautious in the interpretation of the results. First, the large presence of administered and controlled prices may have reduced the responsiveness of consumer prices in the past. However, as prices are progressively liberalized, the pass-through coefficient is bound to increase. Second, the variability of the exchange rate over the sample period is extremely low. This, by itself, makes it difficult to identify statistically significant relations with other variables. More importantly, though, the pass-through cannot be expected to remain the same under different conditions. There is no guarantee that, when the exchange movements become more pronounced, the economy will react in the same way. If the transition to a new regime is not clearly understood by the public, larger exchange rate movements could easily destabilize expectations.

All in all, these results would seem to encourage a gradual shift away from the past policy of strict exchange rate targeting. Despite the cautionary notes outlined above, the findings confirm that Croatian is far from being a fully dollarized economy, and price indexation to the exchange rate has been limited in the past. Since 1994, through a history of low inflation, the CNB has established strong credentials and the new central bank law has now fully recognized its independence and primary focus on inflation. As supervisory oversight improves and prudential regulations are adjusted to ensure that currency risk is fully taken into account by banks, the financial system is becoming less vulnerable. The development of financial markets provides residents with instruments to hedge against currency risk. As a consequence, balance sheet effects are bound to become less important, strengthening the impact of monetary policy.

However, the findings of these paper cannot be interpreted as supporting a policy change. More research is needed to fully evaluate the alternatives facing the policy makers. The past successes of monetary policy in reducing inflation and the persistent vulnerability of the financial system to exchange rate movements argue for a conservative approach.

A. Figures

Figure A1. Cross-Correlations Between the Exchange Rate and the RPI (Upper Panel) and the MPI (Lower Panel)

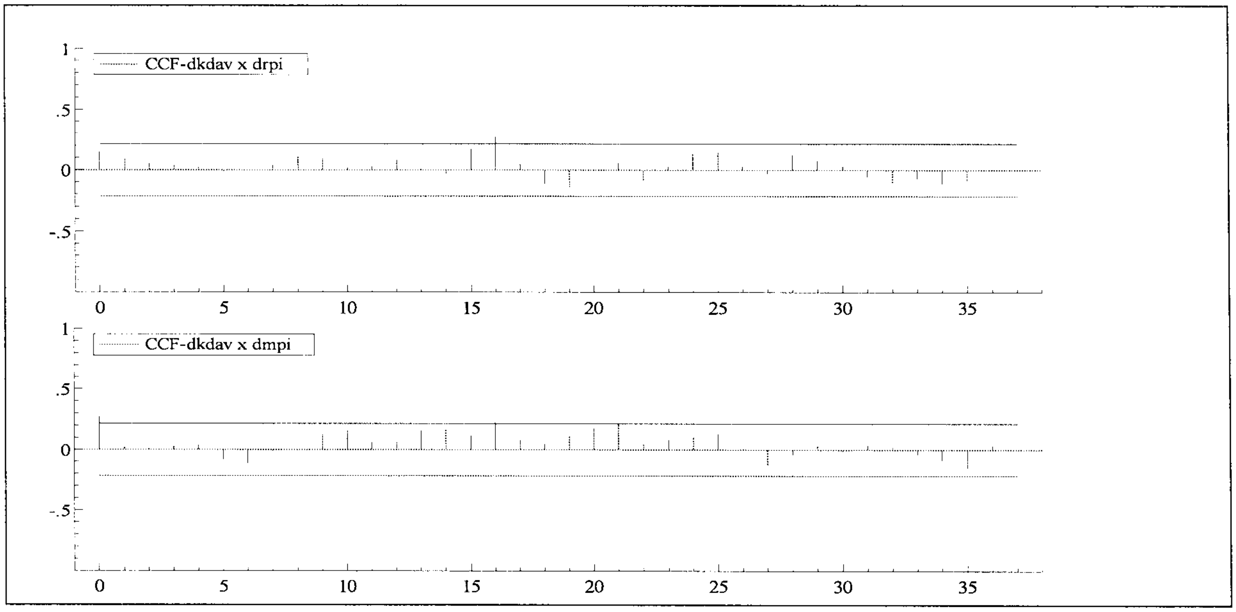


Figure A2. Industrial Production and the Output Gap

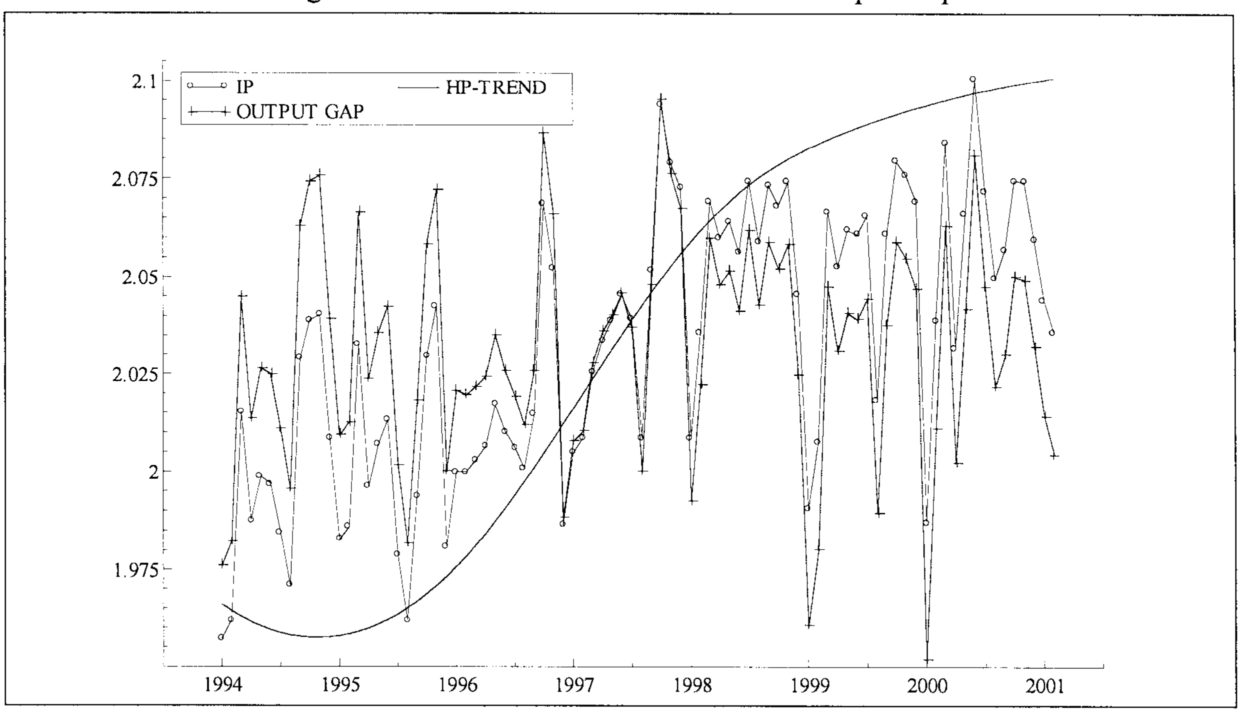


Figure A3. VAR Residuals

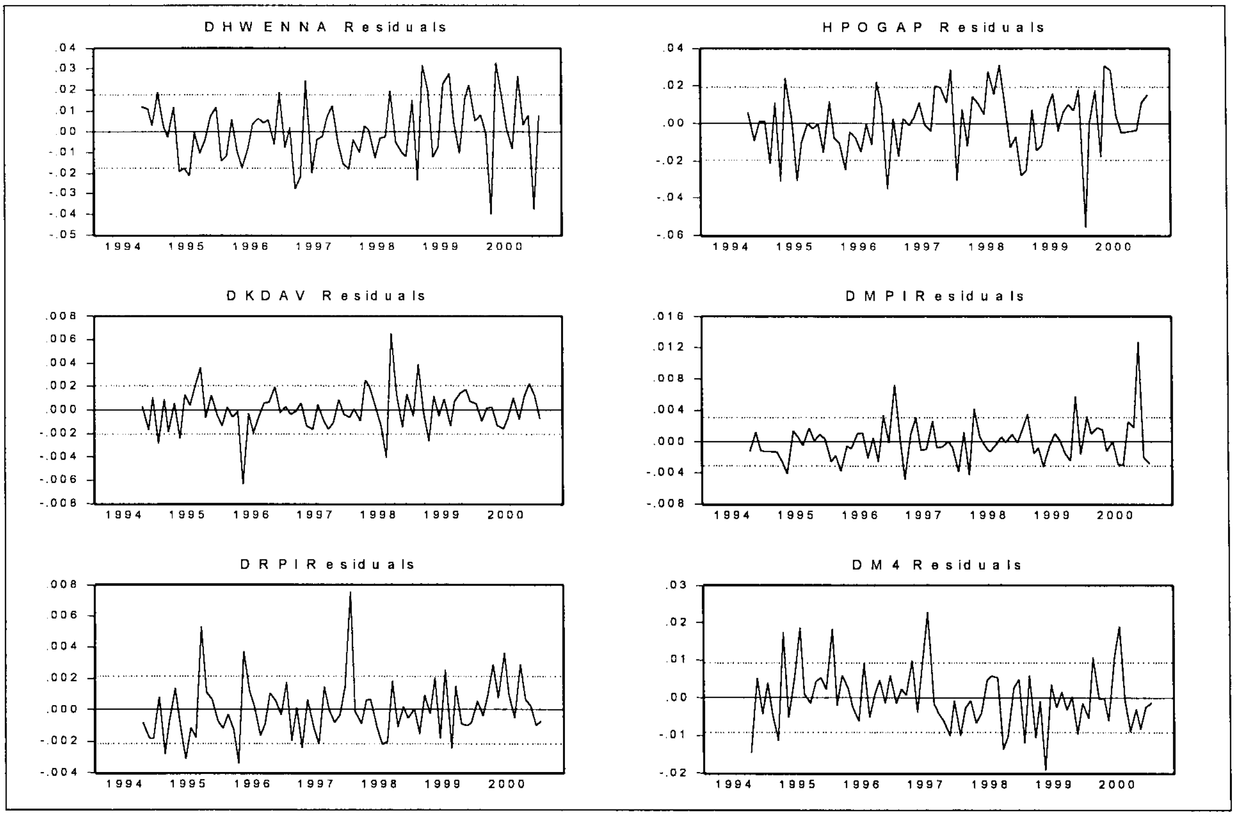


Figure A4. Variance Decomposition for the Change in the HRK-DEM Exchange Rate

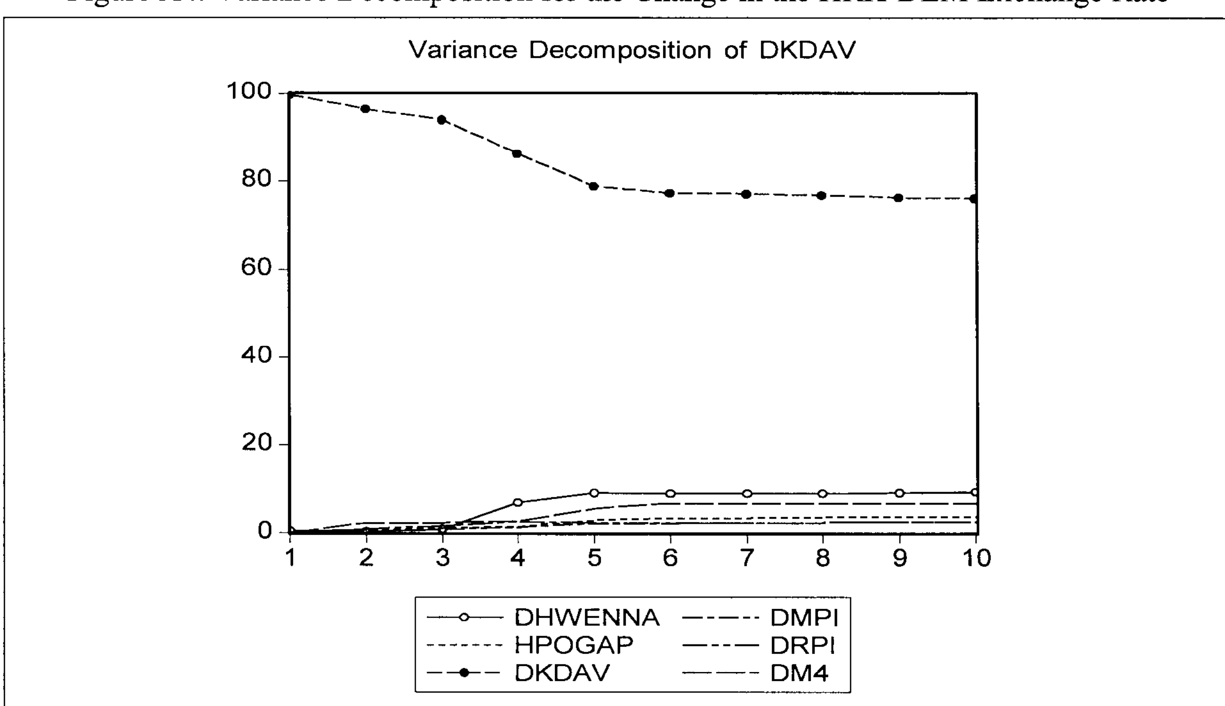


Figure A5. Variance Decomposition for the Change in the Manufacturing Price Index

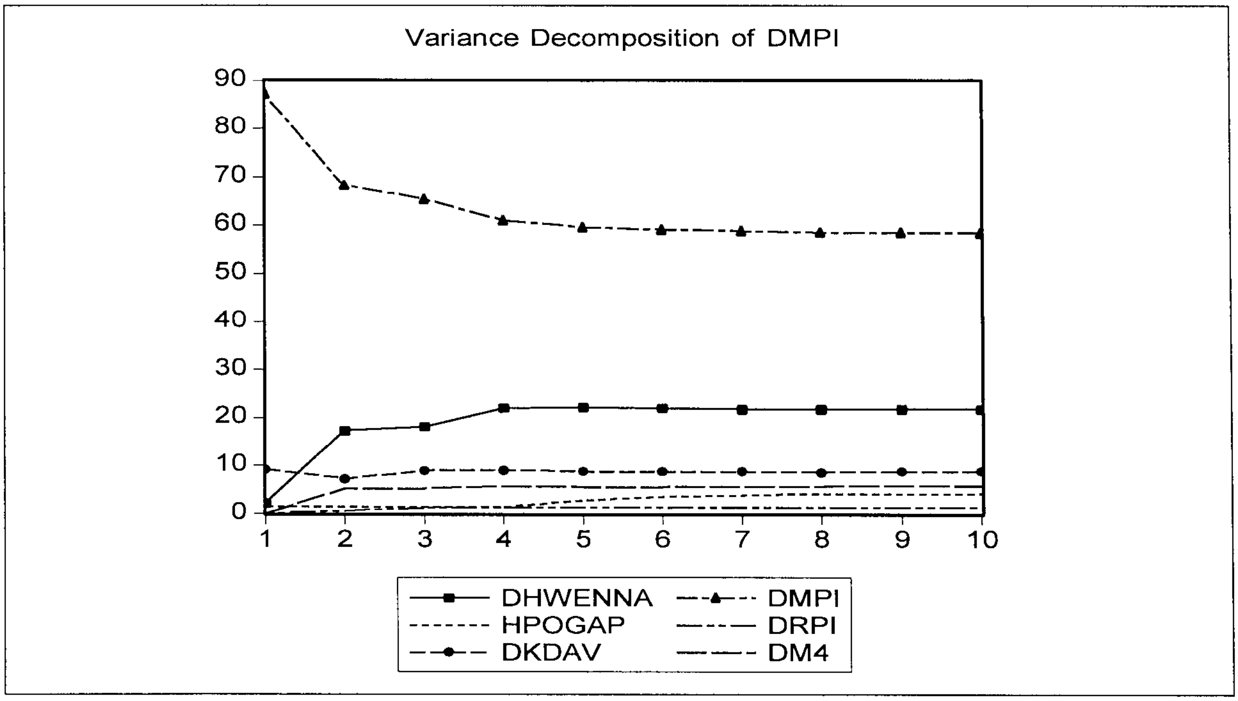


Figure A6. Variance Decomposition for the Change in the Retail Price Index

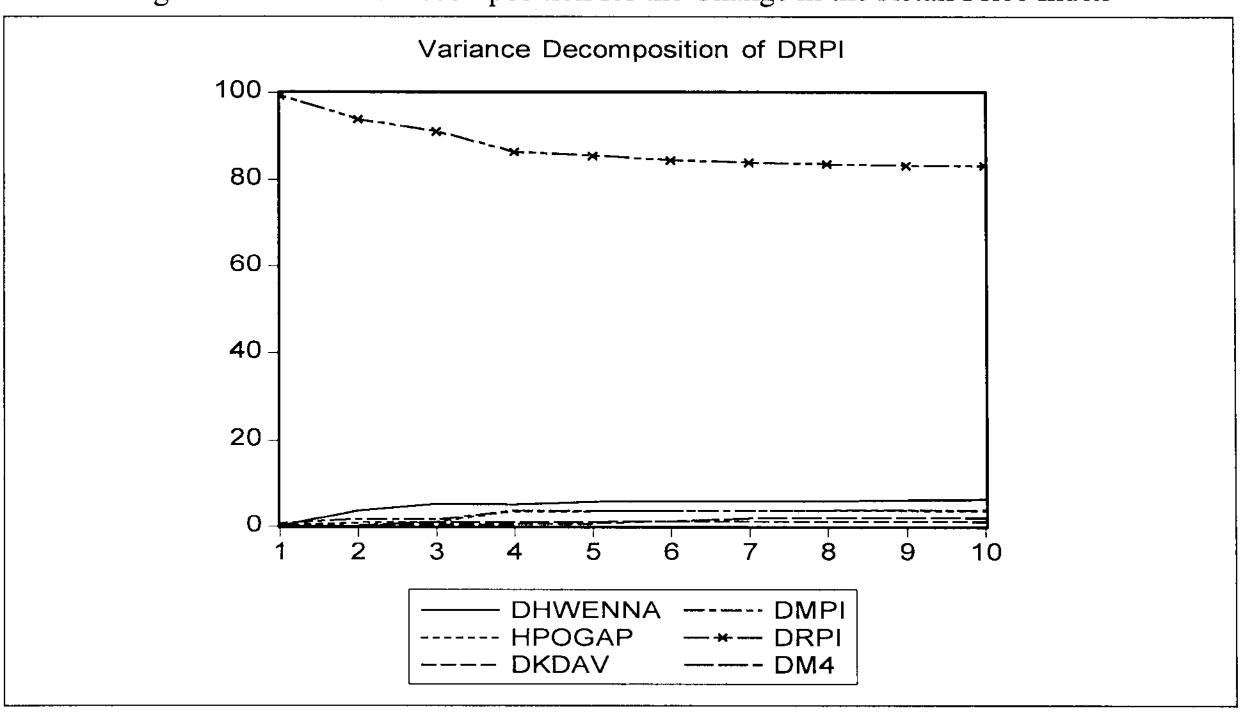
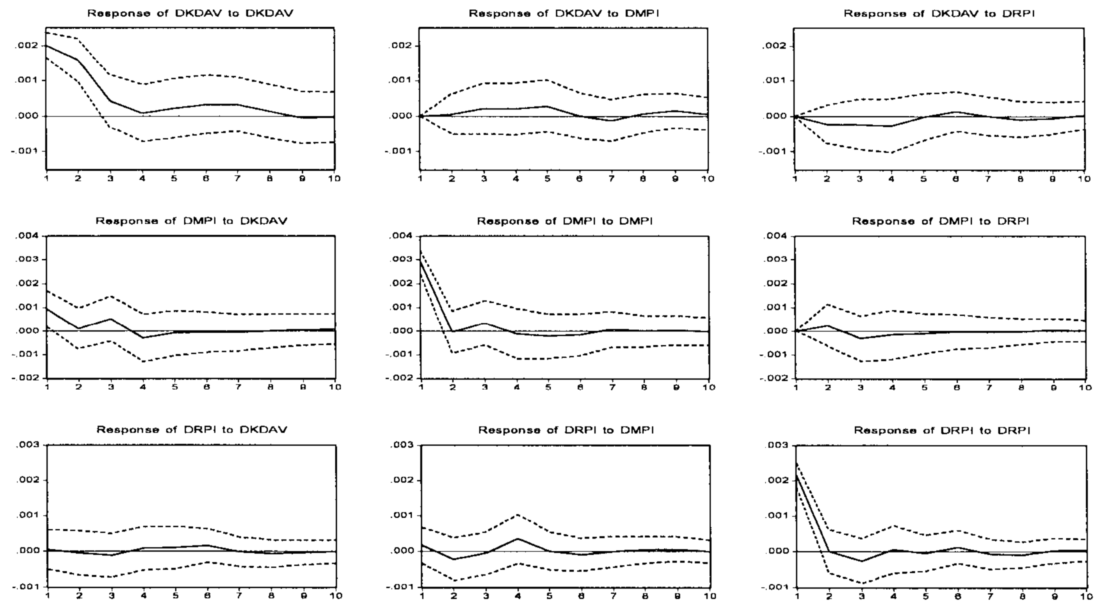


Figure A7. Impulse Response Function for the VAR in First Differences (10 Periods).

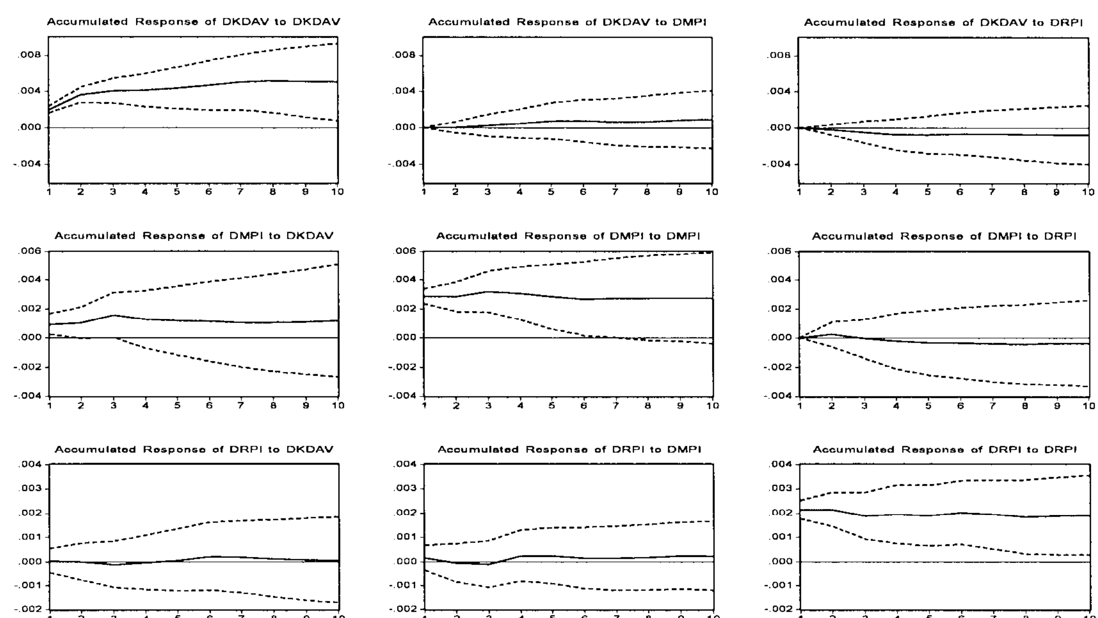
Response to Cholesky One S.D. Innovations ± 2 S.E.



Note: Dotted lines reflect two standard error bands, calculated by Monte-Carlo simulation, 1000 repetitions.

Figure A8. Accumulated Impulse Response Functions (10 periods)

Accumulated Response to Cholesky One S.D. Innovations ± 2 S.E.



Note: Dotted lines reflect two standard error bands, calculated by Monte-Carlo simulation, 1000 repetitions.

B. Tables

Table B1. Pairwise Granger Causality Tests on First Differenced Time Series (no Estimation)

	Lag length in months					
	1	2	3	4	5	6
DHWWA						
DHWWA→HPOGAP	1.72	0.56	0.73	0.73	0.75	0.68
DHWWA→DKDAV	1.09	0.41	0.64	1.02	1.96*	1.52
DHWWA→DMPI	2.92*	3.14**	3.70**	3.16**	3.05**	2.73**
DHWWA→DRPI	0.85	1.45	1.20	1.72	1.10	0.89
DHWWA→DM4	0.22	0.73	1.83	1.79	1.56	1.14
HPOGAP						
HPOGAP→DHWWA	4.65**	2.55*	3.87**	3.22**	2.95**	2.41**
HPOGAP→DKDAV	0.89	0.23	0.50	0.76	2.34**	2.57**
HPOGAP→DMPI	1.93	1.04	0.79	0.34	0.24	0.25
HPOGAP→DRPI	0.61	0.62	2.05	1.60	2.90**	3.94***
HPOGAP→DM4	2.41	1.71	1.01	0.79	0.83	2.16*
DKDAV						
DKDAV→DHWWA	0.05	0.61	0.38	0.30	0.24	0.34
DKDAV→HPOGAP	1.04	1.56	1.01	1.40	1.11	1.25
DKDAV→DMPI	0.12	0.61	3.79**	0.51	0.29	0.34
DKDAV→DRPI	0.59	0.76	2.03	0.71	1.17	0.38
DKDAV→DM4	4.48**	3.69**	2.29*	1.70	1.26	0.84
DMPI						
DMPI→DHWWA	5.17**	2.95*	2.26*	1.46	2.84**	2.43**
DMPI→HPOGAP	0.02	0.47	0.86	0.94	0.74	1.02
DMPI→DKDAV	0.35	0.73	2.70*	1.86	0.56	0.35
DMPI→DRPI	1.20	3.58**	3.39**	1.59	2.76**	1.05
DMPI→DM4	0.39	0.14	0.29	0.23	0.59	1.09
DRPI						
DRPI→DHWWA	5.20**	2.99*	1.88	1.47	1.43	1.30
DRPI→HPOGAP	0.29	2.86*	4.34***	3.50**	2.67**	2.80**
DRPI→DKDAV	0.23	0.06	10.40***	3.32**	1.72	1.07
DRPI→DMPI	0.47	0.03	2.81**	0.32	0.13	0.19
DRPI→DM4	0.01	0.13	0.32	0.41	0.40	0.57
DM4						
DM4→DHWWA	0.13	0.28	0.22	1.22	1.22	1.74
DM4→HPOGAP	3.65*	1.60	7.26***	5.56***	4.38***	4.43***
DM4→DKDAV	0.00	0.69	2.16	2.25*	1.49	2.54**
DM4→DMPI	0.66	0.63	0.42	0.43	0.41	0.43
DM4→DRPI	0.02	0.06	0.16	0.42	0.42	0.35

Note: Numbers are test statistics of standard F-test for the null hypothesis of "no Granger causality from first to second variable". 1, 2, 3 stars indicate rejection at 10, 5, 1 percent significance level. Variables are in first differences, apart from the output gap (HPOGAP). Observations used for estimation vary due to data availability.

Table B2. Unit Root Tests

	Levels		first differences		Critical values
time series	ADF	PP	ADF	PP	ADF/PP
HWENNA	-2.13	-1.78	-3.61**	-8.27***	1%:-4.07, 5%:-3.46, 10%: -3.16
OIL	-1.88	-1.64	-4.29***	-8.96***	1%: -4.06, 5%:-3.46, 10%:-3.16
IP	-5.03***	-6.25***			1%:-4.07, 5%:-3.46, 10%: -3.16
HPOGAP	-5.76***	-6.52***			1%:-4.07, 5%:-3.46, 10%: -3.16
KDAV	-2.70	-1.92	-5.82***	-5.73***	1%: -4.06, 5%:-3.46, 10%:-3.16
NEER	-2.75	-2.69	-4.76***	-12.04***	1%:-4.07, 5%:-3.46, 10%: -3.16
MPI	-3.15	-2.88	-5.56***	-7.47***	1%:-4.07, 5%:-3.46, 10%: -3.16
RPI	-4.44***	-4.38***			1%: -4.06, 5%:-3.46, 10%:-3.16
no trend	1.85	2.37	-5.60***	-7.35***	1%:-3.51, 5%: -2.89, 10%: 2.58
M4	-1.24	-1.97	-3.64**	-6.66***	1%: -4.06, 5%:-3.46, 10%:-3.16

Note: All variables in logs; the sample period starting in 1994M1 (M6 for the monetary aggregate) varies slightly, according to data availability. 1,2,3 stars indicate rejection at the 10 percent, 5 percent, 1 percent level of the null hypothesis of a unit root. For ADF, 3 lagged differences were used, i.e., 4 lags. In the estimation, a trend and a constant were included. For the PP test, the Bartlett kernel was truncated after 3 lags, as suggested by the Newey-West automatic truncation lag selection, based on the number of observations. The slight differences in critical values are due to differences with regard to the number of observations.

Table B3. VAR Residual Correlations

	DHWWA	HPOGAP	DKDAV	DMPI	DRPI	DM4
DHWWA	1	0.117	-0.062	0.151	0.039	-0.201
HPOGAP	0.117	1	-0.008	-0.104	0.003	-0.129
DKDAV	-0.062	-0.008	1	0.292	0.020	-0.199
DMPI	0.151	-0.104	0.292	1	0.084	-0.141
DRPI	0.039	0.003	0.020	0.084	1	-0.249
DM4	-0.201	-0.129	-0.199	-0.141	-0.249	1

Table B4. VAR Pairwise Granger Causality (Wald Test)

	Variable to be excluded													
Dependent variable	DHWWA (3)		HPOGAP (3)		DKDAV (3)		DMPI (3)		DRPI (3)		DM4 (3)		All (15)	
DHWWA			**8.47**	**(0.04)**	0.32	(0.96)	**9.04**	**(0.03)**	3.47	(0.32)	0.80	(0.85)	24.67	(0.05)
HPOGAP	1.88	(0.60)			2.66	(0.45)	5.29	(0.15)	4.90	(0.18)	**19.59**	**(0.00)**	**37.59**	**(0.00)**
DKDAV	7.04	(0.07)	1.33	(0.72)			1.25	(0.74)	2.41	(0.49)	6.15	(0.10)	17.94	(0.27)
DMPI	**19.89**	**(0.00)**	2.38	(0.50)	1.58	(0.66)			1.46	(0.69)	4.89	(0.18)	24.52	(0.06)
DRPI	2.93	(0.40)	3.10	(0.38)	0.25	(0.97)	1.59	(0.66)			0.44	(0.93)	9.81	(0.83)
DM4	7.73*	(0.05)	2.36	(0.50)	7.56	(0.06)	4.85	(0.18)	0.79	(0.85)			19.53	(0.19)

Note: Numbers are test statistics of pairwise Wald-type Granger causality tests in the estimated VAR, taking into account all lags of the respective variable. The test under the null of "no Granger causality" follows a χ^2 distribution; the degrees of freedom are given in parenthesis. Sigificant test statistics (at the 5 percent level) are in bold.

C. Data Sources

Supply side shocks are approximated by the HWWA index of raw materials, compiled by the Hamburg Institute for Economic Research (HWWA), described in Matthies and Timm (1997), and provided through DataStream.

Demand side shocks are approximated by the output gap, constructed as the difference between the IFS time series for industrial production (row 66) and a Hodrick-Prescott-filtered trend.

The exchange rate is the monthly average of the midpoint exchange rate, Croatian kuna (HRK) per 1 deutsche mark, (DEM), provided by the Croatian National Bank.

The manufacturing price index and the retail price index used in the study are produced by the Croatian Central Bureau of Statistics.

The monetary aggregate M4 measures end of period broadest money and is compiled by the Croatian National Bank.

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