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Cointegration Analysis in an Inflationary Environment: What Can We Learn from Ukraine's Nominal Exports?

by

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Cointegration Analysis in an Inflationary Environment: What Can We Learn from Ukraine's Nominal Exports?*

Abstract:

Ukrainian exports can be explained by standard demand theory in the long run. Using the Johansen procedure the data do not reject the hypothesis of a unit foreign-production elasticity of Ukrainian exports, which are rather price-elastic inputs for foreign producers. It is argued that due to high domestic inflation and substantial real appreciation of the hryvnia there might be a deterministic element in the long-run relationships. When allowing for a trend in the cointegration space, the identifying restriction of an infinitely price-elastic export supply curve produces best results. However, due to missing export price statistics long-run interpretations are to be taken with care because they are conditional upon assumptions on how costs and exchange-rates are passed through on export prices.

Keywords: cointegration analysis; transitional economies; Ukraine; export demand; foreign trade elasticities; real effective exchange rate

JEL Classification: F17; F31; F41; P33

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Contents

1. Motivation

The breakdown of former soviet centrally planned economic systems more than ten years ago led to temporary contractions in overall economic activity as the market-price driven reallocation of resources made obsolete large parts of the existing capital stock while new investment needed time to be decided on. The transformation crisis has been particularly deep and prolonged in Ukraine where real per-capita GDP decreased at an average rate of 8.1 percent per year from 1990 to 1999. A huge and growing literature has emerged that aims at explaining the reasons for disparities in the success of transformation. Among the determinants of successful transition most commonly emphasized are a swift implementation and rapid general acceptance of institutions of the market economy (Havrylyshyn, Izvorski and Van Rooden 1999) but also the conduct of sound industry and macroeconomic policies (Havrylyshyn and Van Rooden 2000). The reorientation of trade flows is seen as a source of welfare, too, not only because it has started to reshape the international division of labor according to comparative advantages rather than former political preferences but also because there is now a huge amount of trade between countries formerly separated by the iron curtain. Many empirical studies therefore focus on the progress made so far in reorientation of trade and resource flows using an array of indicators and estimation techniques. The gravity model is the established work horse of this type of analysis and is often used (e.g. Piazolo 2001, Buch and Piazolo 2001 for EU accession states; Kulpinsky 2001 for Ukraine).[1] However, econometric analyses particularly focusing on Ukrainian foreign trade problems are rare in the literature.

While gravity models deal with the regional composition of a country’s exports and imports, a question often asked in applied international economics hinges on

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[1] See Deardorff (1998) for the derivation of gravity models from standard trade theory.

the development of aggregate exports over time and the relative importance of its determinants. Yet time series regressions are absent from the empirical literature on Ukraine's foreign trade to my knowledge.[2] One reason is that all available time series on foreign trade only start in 1993I. I am convinced that with 32 observations preliminary, although not most reliable, conclusions on the empirical influence of export determinants can be drawn and that it is time to explore the macroeconomic reasons of Ukraine's poor export performance econometrically. The second reason for the lacking interest in time series analysis of Ukrainian exports could be that the non availability of export prices makes an investigation into export demand and supply little attractive at first sight. However, it is shown in this paper that cautious conclusions on the relevant long-run relationships are possible although direct answers can only be given for nominal exports.

The question how changes in foreign economic activity and in price competitiveness influence aggregate exports attracts substantial interest because the real exchange rate is one important determinant of exports. Policy-makers have often sought to boost aggregate production and employment via exports by manipulating the exchange rate of the domestic currency. In the empirical part of the paper, long-run elasticities are singled out using the Johansen procedure to Ukraine's aggregate exports, an index of foreign industrial production and the real effective exchange rate. In the major part of the paper (except of section 6.3.) I focus on the demand side assuming an infinitely price-elastic long-run supply curve, which is in line with the major part of the empirical literature (Goldstein and Khan 1985: 1087; Sawyer and Sprinkle 1999: 10).

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[2] In the English-speaking literature time series regression techniques focusing on Ukraine are used to assess the degree of price convergence within the country (Conway 1999) and between Ukraine and the United States (Cushman et al. 2001).

The reminder of the paper is organized as follows. First a production function with constant elasticities of substitution (CES) is presented from which demand for real exports can be derived theoretically (section 2). This framework is then modified to deal with the non-availability of real export figures for Ukraine (section 3). Using the simplifying assumption that foreign production does not influence the price level of exports in the long run, it is argued that the real export demand framework still holds but that the measured long-run elasticity of (nominal) exports with respect to the real exchange rate has to be interpreted with caution (section 4). In section 5 the peculiarity of an inflationary environment and its implications for a plausible exchange-rate pass-through are discussed. Finally a richer export system for Ukraine is estimated for the period 1993I to 2000IV by allowing for a trend in the cointegration space and using constant returns to scale in foreign production as well as a horizontal export supply curve as the preferred identifying restrictions for the two vectors found (section 6). Section 7 summarizes and draws some policy conclusions.

2. The demand for Ukrainian exports

Let foreign firms produce goods and services combining a bundle of their own (foreign) factors (H*) and Ukrainian goods and services (X) using a technology characterized by constant elasticity of substitution (CES):[3]

[1] $$Y^* = \left(a_1 H^{*-g} + a_2 X^{-g}\right)^{-\frac{j}{g}}$$

where j is the scale elasticity ($j = 1$ for constant returns to scale) and $[-1/(1+\gamma)]$ is the elasticity of substitution between H* and X, a_1 and a_2 are the income

[3] The following analysis is inspired by Sandermann (1975: 41 ff.). Clostermann (1998: 204 f.) applies production theory to derive the demand for German exports.

shares of the factors of production. An asterisk symbolizes foreign variables. The foreign firm maximizes its profits (revenues less costs) according to

[2] $$p^*\left(H^*, X\right) = P^* Y^* - P_{H*} H^* - P_X \cdot W \cdot X \ ,$$

where P* is the price level of foreign output, P_{H*} is the price of one unit of the foreign factor and $P_X \cdot W$ is the price for one unit of Ukrainian exports in foreign currency (*W* being the nominal exchange rate in hryvnia per unit of the foreign currency). Substituting [1] into [2], deriving the first-order condition with respect to exports, then using $a_1 H^{*^{-g}} + a_2 X^{-g} = Y *^{-\frac{g}{j}}$ from [1], taking logarithms (symbolized by small letters, i.e. x = lnX) and solving for x yields

[3] $x = h_0 + h_1 y^* + h_2 e$,

where h_0= [1/(1+*g*)] ln(*j* a_2); h_1 = (*j*+*g*)/[*j* (1+*g*)] = h_2(*j*+*g*)/*j* and h_2= 1/(1+*g*); and *e* = *w*+*p**-p_x. *e* is the logarithm of the real effective exchange rate of the hryvnia and increases in *e* and *w* correspond to real and nominal depreciations of the Ukrainian currency, respectively. [3] is the demand for Ukrainian exports expressed in constant domestic prices.

When production in the foreign economy occurs at constant returns to scale, h_1= 1 whatever elasticity of substitution prevails. In the presence of increasing returns to scale (*j* >1) the elasticity of Ukrainian exports with respect to foreign production ($h_?$) is above 1 only if demand for Ukrainian exports is price elastic (–1< *g* <0)? A low price elasticity (*g* > 0, i.e. 0 < h_2< 1) and a production elasticity $h_?$ above 1 can simultaneously be observed only if returns to scale are decreasing. Whenever the aggregate real-exchange-rate elasticity is empirically found to be comprised between 0 and 1 one would therefore expect a long-run elasticity of exports with respect to foreign production of one or slightly below (allowing for constant or increasing returns to scale). The assumption of constant returns to scale at the level of world production ($h_1 = 1$) will serve as a testable restriction in the cointegration analysis below.

3. Data availability and the construction of regressors

In this section I construct the empirical counterparts of the two determinants of real exports just derived for the period quarterly data are available for (1993–2000). Aggregate foreign production is proxied by a weighted average of the index of industrial production in Ukraine's most important trading partner countries[4] with the share of each of them in Ukraine's merchandise exports serving as weights. To deal with the changing regional pattern of exports (Table 1), the average share of each country from 1996 to 1998 rather than data from a single observation period is taken.

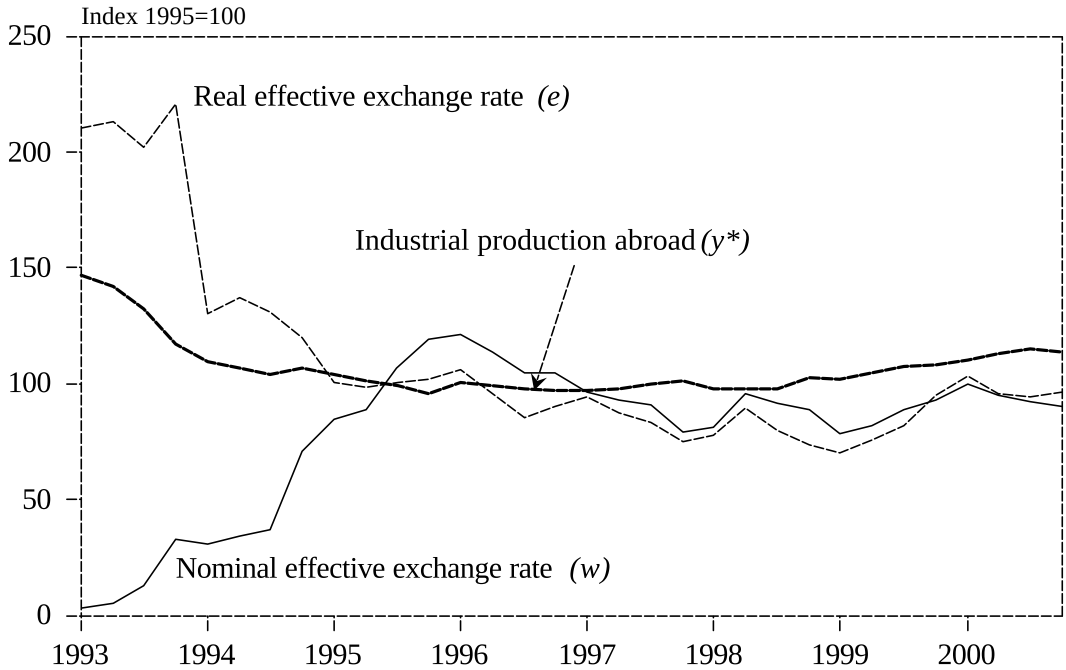
Table 1: Geography of Ukrainian Merchandise Exports 1996–2000 (in percent of total exports)

Year	Russia	Other FSU Countries	EU Countries	Other Countries
1996	38.7	14.6	11.1	35.6
1997	26.2	14.7	12.3	53.2
1998	23.0	12.1	16.8	48.1
1999	20.7	8.8	18.3	52.2
2000	24.1	8.9	16.2	50.8

Source: IMF (2001b); own calculations.

[4] The 25 important countries for which long time series are available are: Poland, Czech Republic, Hungary, Russia, Romania, Slovenia, Germany, France, United Kingdom, Italy, Spain, Netherlands, Belgium, Austria, Sweden, Finland, Denmark, Portugal, Ireland, Greece, Switzerland, Norway United States and Japan. Sources are OECD (2001) and IMF (2001a).

Figure 1: Determinants of Ukrainian Exports[a]



[a]Seasonally adjusted using the multiplicative census-X-11 procedure.

Source: European Commission (2001b); IMF (2001a); OECD (2001); own calculations.

The degree of representation is only 56 percent because for a number of Eastern European countries long time series are not available. Accordingly, the weight of Russia in the sample of trading partners is high (about 53 percent in contrast to an absolute share of 29 percent in Ukrainian exports) and dominates the evolution of the index. The first years shown in Figure 1 exhibit shrinking industrial production reflecting unsettled transformation problems in several CEEC. The timid recovery of 1997 is interrupted by the Russian crisis in 1998 and only the last two years of the sample period show robust economic growth.

The real effective exchange rate of the hryvnia is calculated as the nominal effective exchange rate multiplied by the ratio of foreign to Ukrainian producer prices.[5],[6] First the bilateral real effective exchange rate of the hryvnia is calculated with respect to each partner country, then the real effective exchange rate is constructed as a weighted geometric average of all bilateral real exchange rates. Again the partners' shares in Ukrainian exports serve as a weight.[7] The same geometric weighting is applied to the bilateral nominal exchange rates to derive the nominal effective exchange rate (also shown in Figure 1) which gives an idea of the overall evolution of the Ukrainian currency. For both exchange rate indices the degree of representation amounts to 62 percent.[8] The opposite-signed slopes of the curves at the beginning of the observation period show that the hryvnia has devalued by far less than what would have been implied by the purchasing power parity for tradeables.

Whereas sound empirical counterparts of the theoretical explanatory variables of real exports can be computed, real exports themselves do not yet exist in Ukrainian statistics. Therefore nominal exports have to be used in the estimation. Several measures of aggregate exports of goods and services are shown in Figure 2, each as an index number (1995=100). The "flattest" graph illustrates the index derived from the original data collected by the National Bank of

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[5] When starting from a production function to derive exports as we do in section 2 the theoretically correct domestic price index would be the one of Ukrainian exports, which does not exist.

[6] Only the real exchange rate of the hryvnia relative to the U.S. dollar exists as a long time-series while official figures for the real *effective* exchange rate are only available beginning in 1998 (Kyyak 2001).

[7] The theoretically correct weighting scheme would take account of third-market effects, i.e. the correct share of a specific partner country is its weight in total goods supply competing with the home country's supply on all foreign markets (Deutsche Bundesbank 1998). This computation is impossible due to the lack of a bilateral trade and domestic supplies matrix for the 25 countries involved.

[8] Besides the countries mentioned above, data for Belarus, Latvia, Lituania, the Slovak Republic, and Kyrgystan are available.

Ukraine (NBU) and published in European Commission (2001b: 26), where exports of goods and services are denominated in US dollars. The “steepest” graph expresses the same exports in hryvnia. This series is obtained by multiplying exports in US dollars with the hryvnia-dollar exchange rate; theoretically it corresponds to adding p_x to both sides of equation [3]. If one additionally subtracts w, one gets the third graph showing Ukraine’s nominal exports in units of a representative foreign currency. This is the “export bill” facing the representative foreign firm and therefore is probably the closest proxy for demand for Ukrainian exports.[9] I use this expression of nominal exports in the cointegration analysis of section 4.

The only modifications made so far to [3] thus consist of adding foreign currency prices of Ukrainian exports $(p_x - w)$ to both sides of the equation to make it a structural demand relationship for nominal exports, labeled s in the remainder of the paper:

[4] $s \equiv (p_x - w) + x = h_0 + h_1 y^* + h_2 e + (p_x - w)$

This simple augmentation, however, involves practical estimation problems. Firstly, the unobservability of export prices rules out the possibility of picking the identity $p_x - w = p_x - w$ as a second cointegrating vector to isolate real export demand (as in [3]) from the system. Secondly, the implicite presence of both export and producer prices would potentially inflate the cointegration space, as the relative profitability of exporting is generally seen as the main supply-side

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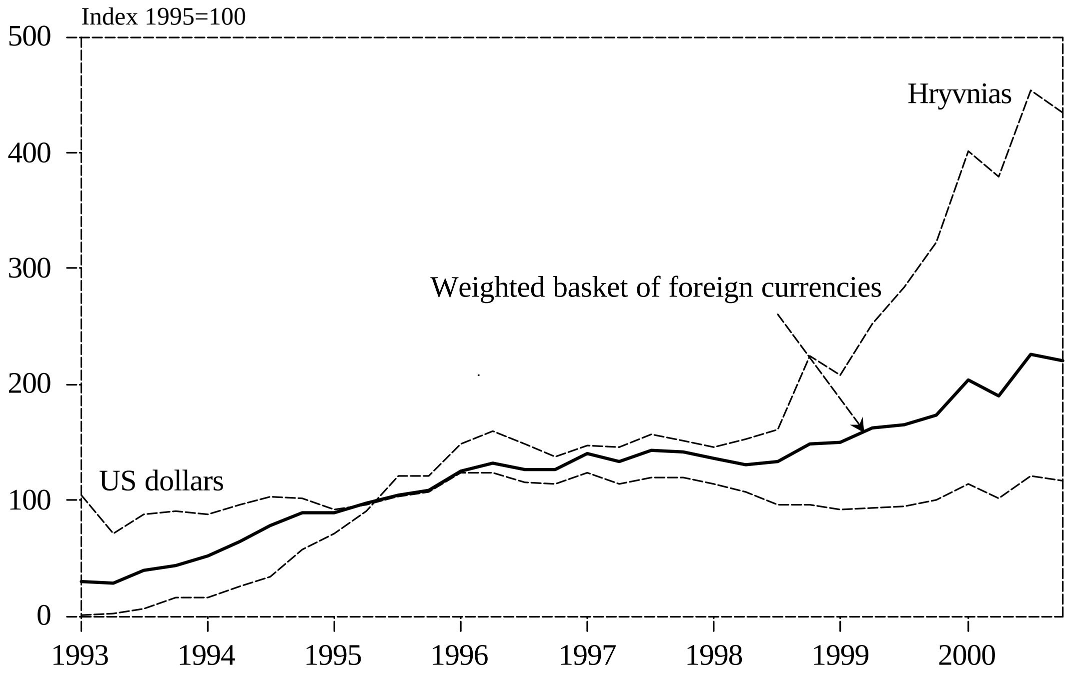
9 According to information by the NBU the lion’s share of both Ukraine’s exports and Ukraine’s imports is invoiced in U.S. dollars. Nonetheless the relative affordability of Ukrainian goods is better expressed by prices in an export-weighted basket of currencies as customers have to buy or earn U.S. dollars before buying Ukrainian goods. As the overwhelming majority of trading partners are non-dollar countries, the price in units of their own respective currency (Russian or Belorussian roubles, euro etc.) ultimately decides on the quantities demanded.

variable. According this view export prices depend either on domestic production costs alone (simple mark-up model) or on both domestic and foreign production costs as is the case in the "Extended Dixit-Stiglitz model" (Dornbusch 1987: 99-100). Export prices might therefore be cointegrated with the real exchange rate. In this case the "true" value of h_2 could not be estimated because the empirical coefficient would result from a linear combination of demand and supply influences. The rank test in the following section serves to detect the number of cointegrating vectors and thus to see if such an ambiguity problem exists in the problem at hand. As supply-side variables cannot enter the system directly, the sensitivity of rank test results to justifiable modifications of specification is given ample room in the following sections.

Figure 2: Ukrainian Nominal Exports, Expressed in Different Currencies[a]



[a]Seasonally adjusted using the multiplicative census-X-11 procedure.

Source: European Commission (2001b); IMF (2001a); own calculations.

4. Cointegration Analysis I: Identifying the Demand Relationship

4.1. Johansen Procedure

An approach particularly suited to verify theoretical long-term equilibrium relationships in the presence of non-stationary time series[10] is the procedure developed by Johansen (1991) as well as Johansen and Juselius (1994). The series used are: the log of the index of Ukraine's nominal exports of goods and services in "average" units of foreign currency (denoted s, where $s = x+p_x–w$), the log of the weighted index of industrial production abroad (y^*) and the log of the real effective exchange rate of the hryvnia (e). The analysis starts with the vector error correction model (VECM)

[5] $$\Delta z_t = \prod z_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta z_{t-i} + m + u_t \ ,$$

where z represents the vector of non-stationary variables s, y^* and e; m denotes the vector of the constant terms, u the vector of the iid residuals. The rank of matrix P indicates the number of cointegration relationships in the system. If P is of reduced rank ($0 < r < p$), it can be separated ($P = ab'$) into a (p x r)-dimensional matrix of loading coefficients a and a (p x r)-dimensional matrix of the cointegration vectors b, which represent the long-run economic relationships. Our aim is to detect the demand relationship derived in [3]. The lag length k of the model is chosen to be 2 according to the Akaike information criterion.[11]

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[10] The time series used are integrated of order one (see the unit root tests in the appendix).

[11] The AIC reaches absolute minima of equal size (–18.23) for $k = 2$ and $k = 4$. Given the small number of observations $k = 2$ is preferable. The Schwartz criterion (SC) favors $k = 1$, which would eliminate the whole short-run dynamics apart from the intercept. However, residuals are autocorrelated of order 1 if $k = 1$, so minimizing SC subject to well-behaved residuals also leads to $k = 2$.

Another important preliminary choice regards the deterministic components of the model. Whenever the time series used are consistently upward or downward sloping over time (as Ukrainian nominal exports), it is advised to incorporate an unrestricted constant into the model in order to capture systematic influences potentially not explained by the I(1) variables (Hansen and Juselius 1995: 6). I therefore allow for an unrestricted constant in the system.[12]

The number of cointegration vectors is then determined by the rank test.[13] Table 2 summarizes the results. The null hypothesis of only one cointegration relationship ($r = 1$) cannot be rejected. Specifically, the coexistence of an export demand and a long-term relationship between the exchange rate and the level of export prices is not corroborated in the specification chosen. It therefore is suitable to restrict the VECM to one cointegration vector.

Table 2: Test for Co-integration Rank in the VECM (k = 2, unrestricted constant)

Null hypothesis	Trace statistic	Adjusted trace statistic[a]	Critical values (90 %)[b]
r < 1	67.78	54.22	26.70
r < 2	16.48	13.18	13.31
r < 3	0.00	0.00	2.71

[a]Using Reimers' (1992) adjustment; the ratio between the adjusted trace statistic and the conventional trace statistic is $\frac{T - pk}{T}$ where T is the number of observations (30), p the number of I(1) variables in levels (3), and k the lag length (2). — [b]As reported in Hansen and Juselius (1995: 81, Table B.3).

12 In addition to that, a linear trend restricted to the long-run relationship might be considered as well, especially if the omitted systematic growth or decline economically belongs to the long-run relationship. Such a specification is discussed in section 6.

13 The subsequent analysis uses the procedure CATS in RATS (cf. Hansen and Juselius 1995).

4.2. The System Restricted to One Cointegration Vector

To see whether the vector found is an export demand relationship I additionally restrict the long-run production elasticity to unity. The restriction is not rejected by the data ($c^2 = 1.25; \mathit{probability}[0.26]$). The valid constant-returns-to scale

Table 3: Results for the Ukrainian export demand system[a, b, c, d]

[1] $$\Delta s_t = -\underset{(-4.06)}{0.53} - \underset{(-4.57)}{0.08}\left[s_{t-1} - \left(y^*_{t-1} + \underset{(5.64)}{1.77} e_{t-1}\right)\right] - 0.34\Delta s_{t-1} + \hat{u}_{1t}$$

(STDDEV $= 0.062$; $R^2 = 0.55$; normality $= 1.20$; Arch(2) $= 0.92$)

[2] $$\Delta y^*_t = \underset{(4.16)}{0.20} + \underset{(4.10)}{0.03}\left[s_{t-1} - \left(y^*_{t-1} + \underset{(5.64)}{1.77} e_{t-1}\right)\right] + \hat{u}_{2t}$$

(STDDEV $= 0.023$; $R^2 = 0.58$; normality $= 3.12$; Arch(2) $= 1.66$)

[3] $$\Delta e_t = \underset{(1.32)}{0.31} + \underset{(1.39)}{0.04}\left[s_{t-1} - \left(y^*_{t-1} + \underset{(5.64)}{1.77} e_{t-1}\right)\right] + \hat{u}_{3t}$$

(STDDEV $= 0.111$; $R^2 = 0.18$; normality $= 5.71$*; Arch(2) $= 2.50$)

Multivariate statistics[e]:

$L - B(8) = 58.26[0.23], LM(1) = 17.64[0.04], LM(4) = 8.19[0.51],$

normality : $c^2(6) = 11.33[0.08]$

* (**, ***) denotes rejection of the null at the 10 (5, 1) percent significance level.

[a]The initial VECM is restricted to r=1 (one cointegrating vector). The long-run elasticity of Ukrainian exports with respect to foreign production is restricted to unity implying constant returns to scale in foreign industrial production. — [b]t-values in brackets. — [c]Apart from the intercept only significant short-run coefficients are reported. — [d]Univariate statistics are reported below each equation: the standard deviation of residuals, the R² of the equation, the univariate Doornik and Hansen (1994) statistic for normality and the statistics of the Arch test of order 2 for heteroskedasticity. — [e]The Ljung-Box and Lagrange Multiplier tests for autocorrelation and the multivariate test for normality by Doornik and Hansen (1994). The figures in square brackets are probabilities.

restriction means that a one percent increase in foreign industrial activity leads to a one percent increase in Ukrainian exports. The cointegration results are subject to this restriction (Table 3).

Turning to the real effective exchange rate, a one percent real depreciation of the hryvnia (*e* rises) triggers an increase in nominal exports by 1.77 percent suggesting foreign demand for an average basket of Ukrainian commodities is rather price-elastic. It seems that Ukraine's offer on markets for metals, food items, non energy raw materials and chemicals, which together represent almost two thirds of total merchandise exports, can be substituted by goods from elsewhere quite easily; in contrast, fuel and energy products which are known to be price-inelastic in demand only represent a minor part in the country's export receipts (Table 4).

Table 4: Commodity Structure of Ukrainian Exports 1996–1999 (in percent)

Category	1996	1997	1998	1999
Fuel and energy products	7.9	8.1	7.5	8.5
Food items and raw material	19.6	7.7	10.1	11.4
Wood and wood products	1.2	1.7	1.7	2.5
Chemicals	14.1	13.7	12.7	11.1
Ferrous and nonferrous materials	30.0	39.3	39.0	39.1
Industrial products	3.9	4.5	4.9	4.8
Machinery	13.3	13.1	13.0	11.1
Other	10.0	12.0	11.2	11.5
Total	100	100	100	100
Memorandum Item: Total Exports in mill. of US dollars	15 547	15 418	13 699	12 463

Source: IMF (2001b); own calculations.

The long-run export demand relationship enters the equations for both exports and foreign production with significant, correctly-signed but rather small loading coefficients. If exports overshoot their long-run level given by the levels of y^* and e, they are to shrink in the quarters afterwards thus gradually correcting their initial disequilibrium. The adjustment process is supported by a slight uptick in foreign activity as trading partners receive more Ukrainian inputs than would be optimal at prevailing prices. Yet the new export level can be at least partly rationalized by higher price competitiveness. Thus it is straightforward that overshooting exports are followed to a slight real depreciation due either to a trimming in Ukrainian producer prices or to a nominal depreciation. Still, the positive loading in equation [3] of Table 3 is statistically insignificant. Its *t*-value is small (1.39) and the hypothesis of e being weakly exogenous cannot be rejected.[14] However the real exchange rate is maintained as an endogenous variable because setting the loading to zero would lead to violation of the requirement of normally distributed residuals.

4.3. Analysis of a partial system

From an economic standpoint it would be more satisfactory to find foreign production being exogenous (which is not the case here) as Ukraine is a relatively small country in the international context and rather unlikely to influence the Russian, German or American business cycles by its exports, whereas changes in the real exchange rate (taking, for instance, the form of higher domestic inflation) may well occur in times of external imbalances. Based on this economic

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[14] Actually, the loading near zero might be the result from conflicting forces. Rather than accommodating prices to the new export level (as described in the text) Ukrainian exporters could as well perceive the high level as temporary and wish to ration the market via price increases. The overshooting in s could also stem from “too” high export *prices* (not volumes) leading to either a currency depreciation or a reduction in export prices. In the latter case, lower export prices should feed back into producer prices to maintain some profitability of exporting activities. This feedback is then observed by a rising e.

"prejudice" statistical rejection of the exogeneity presumption for y^* is now ignored and foreign production is introduced as an exogenous variable into the system. Once again $r = 1$ is set and the constant-returns-to-scale hypothesis imposed.[15]

The introduction of the contemporaneous change in foreign industrial production improves the R^2 of the export equation from 0.55 to 0.69 and lowers the standard deviation of the regression indicating an important short-run link between foreign output and Ukraine's deliveries. The problem with first-order serial correlation indicated in Table 3 has now disappeared. The only remaining problem with residuals is that the null of normally distributed residuals is rejected at the 10 percent (not at the 5 percent) significance level in the real exchange rate equation. So with respect to residuals the results are quite reliable (Table 5).

Table 5: Results for Ukrainian exports with exogenous foreign production[a, b]

[1]
$$\Delta s_t = \underset{(-6.18)}{-\,1.02} - \underset{(-6.59)}{0.09}\left[s_{t-1} - \left(y^*_{t-1} + \underset{(5.38)}{2.63}e_{t-1}\right)\right] \underset{(-2.67)}{-\,0.30}\,\Delta s_{t-1} + \underset{(3.68)}{1.49}\,\Delta y^*_t + \hat{u}_{1t}$$

$$(\text{STDDEV} = 0.052;\ R^2 = 0.69;\ \text{normality} = 0.21;\ \text{Arch}(2) = 0.82)$$

[2]
$$\Delta e_t = \underset{(1.66)}{0.58} + \underset{(1.71)}{0.05}\left[s_{t-1} - \left(y^*_{t-1} + \underset{(5.38)}{2.63}e_{t-1}\right)\right] + \hat{u}_{2t}$$

$$(\text{STDDEV} = 0.109;\ R^2 = 0.21;\ \text{normality} = 5.30{*};\ \text{Arch}(2) = 1.17)$$

Multivariate statistics[a]:

$L-B(8) = 18.32[0.69], LM(1) = 4.74[0.31], LM(4) = 2.00[0.73],$

normality : $c^2(4) = 7.73[0.11]$

* (**, ***) denotes rejection of the null at the 10 (5, 1) percent significance level.
[a]All footnotes of Table 4 also apply to this table. — [b]Foreign industrial production is introduced as an exogenous variable into the system.

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[15] The test for $\hat{b}_{y*} = 1$ yields $c^2 = 0.09$ (probability of [0.77]) and is clearly accepted.

The long-run elasticity of nominal exports is estimated to be even higher (2.63) than in the model with three endogenous variables. However, this elasticity only gives a rough approximation of the economically relevant price elasticity of real demand for Ukrainian exports. The lack of precision stems from two distinct reasons which are discussed in the following. The first one has to do with the uncertain way in which changes in the nominal exchange rate or in Ukrainian production costs are passed through on export prices. The second one relies on the assertion that the real effective exchange rate as it is calculated here might not measure the country's price competitiveness on foreign markets correctly.

5. Uncertainties surrounding the empirical price elasticity of Ukrainian export demand

Although the finding of only one cointegrating vector seems not to allow any other stationary relationship than the export demand vector, one might question this result. Not only does the small number of observations leave all test results, e.g. the finding of $r = 1$, with some uncertainty, but one may also ask whether the standard specification presented above is appropriate in a period of substantial and permanent real appreciation. The alternative is to capture this typical feature of the Ukrainian transition by allowing for a trend in the cointegration space. In this section I discuss the possible links between the unobservable export prices and the real effective hryvnia exchange rate and their consequences for the long-run price elasticity of exports. In the next section, the model setup is modified to allow for these potential links.

Assume a one percent devaluation of the hryvnia at a moment when export prices had reached a profit-maximizing level. The immediate effect is that Ukrainian goods become cheaper from the perspective of foreign buyers who increase their demand subsequently. The optimal reaction of Ukrainian exporters may consist of both price and quantity adjustments involving a second cointegration vector between w on the one hand and p_x (and thus s) on the other hand. In the extreme case export prices in hryvnias rise to the extent to which the currency depreciates thereby leaving the foreigners' import bill unchanged. This extreme case of pricing to markets is especially relevant for countries that invoice their exports in units of a foreign currency and do not have own discretionary pricing power, e.g. because prices are fixed by long-term contracts, because they are determined on the world market or because part of exports are barter trade with fixed terms of exchange.[16]

Moreover, pricing-to-market behavior is likely to prevail in a highly inflationary environment. If the devaluation is a mere reflex to past domestic inflation, p_x will rise to the same extent as w. The same arguably happens in case of a real depreciation due to foreign producer price inflation being even higher than the Ukrainian one. In both cases exporters can at least partly restore their profit margins suppressed by cost-push inflation. In turn, partial or even full exchange rate pass-through (corresponding to unchanged hryvnia prices and declining export prices in units of the average foreign currency) following a nominal depreciation is to be expected in times of more stable prices.[17] As a consequence, when domestic inflation is high, the elasticity of real exports with respect to the real effective exchange rate should converge to the estimated nominal counter-

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[16] According to information given by the NBU, among Ukraine's major trading partners, Belarus is the only country paying a noteworthy part (one sixth) of its imports from Ukraine in hryvnias. As to barter, it is less of a problem in international exchanges than in the domestic economy, but in 1997 around 10 percent of total goods trade were barter (Worldbank 2001: 68). This share has since declined; it was probably higher before 1997 and therefore might have added some inflexibility to the level of export prices during the sample period.

[17] One normally would analyze the question if different sources of real depreciations have different impacts on export prices (or, in our case, on nominal exports) by introducing the variables composing the real effective exchange rate separately into the model. However, this strategy is not viable because of the severe lack of degrees of freedom.

part — 2.63 (Table 5) and 1.77 (Table 3), respectively. It should be even bigger in times of price stability to make up for changes in the "dollar" price of aggregate exports.

To yield plausible assumptions on what the actual development of Ukrainian export prices could have been, a look at the historical changes in known price indices is helpful (Table 6). As actual export prices are not known, I distinguish between two "stylized" cases.

Case 1: exporters do not raise their prices by more than foreign competitors. Given almost equal average changes in foreign producer prices and the nominal effective exchange rate during the observation period, this implies roughly stable export prices in foreign currency units. Then 2.63 is the correct exchange-rate elasticity of *real* exports.

Table 6: Average quarterly percentage changes in key variables

Variable	1993 I–2000 IV	1995 I–2000 IV
Producer price index	27.8	6.2
Foreign producer price index[a]	11.9	6.7
Nominal effective exchange rate[a]	11.6	0.3
Real effective exchange rate[a,b]	–2.5	–0.2
Nominal exports in $	0.4	1.0
Nominal exports in hryvnia	22.0	8.2
Nominal exports in units of a representative f	6.7	4.0
Foreign industrial production[a]	–0.8	0.4

[a]Weighted by the share of each partner country in Ukraine's merchandise exports. — [b] Based on Ukrainian and foreign producer price indices. — [c] Equals nominal exports in hryvnia divided by the nominal effective exchange rate.

Source: European Commission (2001b); IMF (2001b); OECD (2001); own calculations.

Case 2: exporters maintain their profit margins raising prices to the same extent as domestic production costs. This leads to a fall in *e* (real appreciation) on the right-hand side of equation [4] accompanied by an increase in the "dollar" price level of exports. Thus for the resulting fall in nominal exports (2.63 times the real appreciation) to hold, the elasticity of real exports must exceed 2.63.

How realistic is case 2? If Ukraine's producer price inflation had been fully passed through on export prices, the PPI (as a proxy of the price index for tradeables) could be used to compute a series of real exports. Deflating hryvnia export values in this way generates export volumes that literally collapse from 1993I to 1998I falling to 15 percent of their initial level. Although there is little doubt that real exports have not risen during the period mentioned, the dimension of this decrease seems uncredibly large. I carefully conclude that export prices have not risen as fast as domestic producer prices. Does this mean that case 1 is the relevant one? Not necessarily. If domestic PPI inflation is a good proxy for production cost increases in the tradeables sector, rising export prices only to the extent of foreign producer prices would have forced a considerable share of exporting firms out of the market by driving their profit margins into negative territory. To sum up, the average increase in the unobserved hryvnia export prices from 1993 to 2000 likely was steeper than the rise in the export-weighted foreign PPI but more moderate than the rise in the domestic PPI.

If one accepts the assumption just made on export prices, the second caveat surrounding the interpretation of the empirical price elasticity of export demand becomes relevant. It has to do with the way the nominal exchange rate is adjusted to derive the real one. By lack of reasonable alternatives Ukrainian producer prices are used although the theoretical export demand in [3] suggests using Ukrainian *export* prices. If it is true that the latter increased less than the former, the variable e substantially underestimates the country's competitiveness. Given the plausible assumption of an incomplete exchange rate pass-through, the true price elasticity of demand for Ukrainian real exports is higher than the real exchange rate elasticity. All in all two compelling reasons why the estimated values of 2.63 (and 1.77, respectively) are lower bounds of the economically relevant price elasticity of demand can thus be put forward.

6. Cointegration Analysis II: Accounting for "Permanent" Real Appreciation by the Deterministic Part of the Model

The conclusions of the preceding discussion on export prices suggest it may be advisable to account for the huge and permanent real appreciation during the observation period in the deterministic setup of the model, as additional variables cannot be incorporated. Given strong domestic inflation the unobserved price index of Ukrainian exports should have increased by more than the nominal effective hryvnia exchange rate thereby causing a systematic increase in "dollar" export prices. By allowing for a linear trend in the cointegration space this increase can be made explicit. Yet as it seems unlikely that export prices have risen by as much as the domestic PPI, part of the measured real appreciation is systematic and not captured by the foreign trade data in the model. This is what the deterministic trend could alternatively stand for. The two cases require different identifying restrictions. They are both discussed below thereby giving a better understanding of the possible shapes of Ukraine's export supply.

As a huge amount of uncertainty surrounds the time path of the unobservable export prices and as the number of variables, cointegrating vectors and combinations of (over-)identifying restrictions will be higher than before (see section 6.1.), I follow a rather agnostic strategy to find out the best specification. After a series of tests on the b-vectors two final specifications are presented. They contain the demand vector already known and one of the following relationships as a second vector.

"**Supply 1**": Nominal exports have a positive and significant time trend interpreted as a constant in export price inflation to reinforce the assertion that at least part of the consistently positive differential between Ukrainian producer price inflation and the average increase in the nominal effective exchange rate is passed through on foreign customers. In this setup the underlying price elasticity of export supply is not priori restricted to infinity. The identifying restriction chosen for this vector is the assumption that y^* as a typical demand variable has no long-run influence on export prices.

"**Supply 2**": Here a null restriction is put on exports ultimately implying a horizontal supply curve again. Infinitely elastic export supply means export quantities have no long-run influence on export prices and thus on producer prices[18] and thus on the real effective exchange rate.

6.1. Rank Test

To ensure the highest possible comparability of results, k = 2 is chosen. The null hypothesis of at most one cointegrating vector is now clearly rejected in the rank test using Reimer's adjustment while the null one of $r \leq 2$ is not rejected (Table 7).

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[18] In most models of export supply the same unit resource cost is assumed to prevail in tradeables production for domestic and foreign markets (De Grauwe 1988: 64). So as long as domestic and foreign profit margins are cointegrated, producer prices cannot deviate indefinitely from export prices in theory.

Table 7: Test for Cointegration Rank in the VECM (k=2, trend in the cointegration space)

Null hypothesis	Trace statistic	Adjusted trace statistic[a]	Critical values (90 %)[b]
r < 1	84.39	67.51	39.08
r < 2	32.67	26.14	22.95[c]
r < 3	3.66	2.93	10.56

[a]Using Reimers' (1992) adjustment; the ratio between the adjusted trace statistic and the conventional trace statistic is $\frac{T - pk}{T}$ where T is the number of observations (30), p the number of I(1) variables in levels (3), and k the lag length (2). — [b]As reported in Hansen and Juselius (1995: 81, Table B.3). — [c] Use of critical values by MacKinnon et al. (1999) leads to rejection of the null at the 10 percent (23.34) and even the 5 percent (25.86) significance levels, as well.

6.2. Narrowing the Range of Potential Cointegrating Vectors

While both “Supply 1” and “Supply 2” are just identified, I again postulate constant returns to scale in average foreign production to identify real export demand. I additionally set the trend in the demand vector equal to zero in both specifications in order to pick the same demand vector as in the previous section. This over-identifying restriction is accepted by the data (Table 8).

But one also wants to be sure that none of the vectors is a linear combination of relationships involving the other one. Therefore it is tested whether anyone of the relevant cointegration relationships is present in each of the two vectors. This is clearly rejected for the just identified (“Demand just”) and the over-identified demand vector (“Demand over”), for “Supply 1” and “Supply 2”. To further limit potential arbitrariness of specifications I look if the second vector just comes in by a trend-stationary behavior of *s, e* or *y**. This analysis is a multivariate complement to the unit root tests presented in the appendix. For exports and the real effective exchange rate the hypothesis is clearly rejected. The

Table 8: Tests on subsets of the cointegration vectors (*k=2*, *r=2*, trend in the cointegration space)[a]

Hypothesis	c^2-Statistic	DF[b]	Probability
"Demand over" in cointegration space	1.68	1	0.20
"Demand just" in both vectors	19.35	2	0.00
"Demand over" in both vectors	23.48	4	0.00
"Supply 1" in both vectors	22.33	2	0.00
"Supply 2" in both vectors	8.12	2	0.02
s trend-stationary[c]	9.72	1	0.00
y^* trend-stationary[c]	2.34	1	0.13
e trend-stationary[c]	7.43	1	0.01[d]

[a]The methodology follows Hansen and Juselius (1995: 34–44). — [b]Number of degrees of freedom. — [c]The coefficients of the other two I(1)-variables are restricted to zero. — [d]Rejection at the 1 percent level (critical value is 6.64).

null of trend-stationarity in foreign industrial production cannot be rejected at the 10 percent level, although the decision is tight. However, this result is ignored for two reasons. Firstly, it is in sharp contrast to the ADF and Phillips-Perron tests which see y^* rather on a knife's edge between I(1) and I(2) than between I(0) and I(1) (Table A.1). Secondly, and most importantly, the trend stationarity of y^* would contradict the clear rejection of the hypothesis that the demand relation is present in both vectors as "Supply 1" would then just replace the restricted y^* by a coefficient of the time trend.[19]

6.3. Allowing for An Upward-Sloping Export Supply Curve

As by now one may be more comfortable about the soundness of the two final specifications, what can be learnt from the first one which is shown in Table 9? The residuals are well behaved without any need for contemporary first differ-

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[19] Strictly speaking, a demand vector just identified by the absence of a time trend (but with an unrestricted coefficient for y^*) would be observationally equivalent to "Supply 1". The test (not reported in Table 8) for presence of such an alternative "demand" relationship in both vectors is also rejected ($c^2(2) = 12.95[0.00]$).

ences. The explanatory powers of all equations are better than in section 4, even substantially so for the equation of the change in the real effective exchange rate ($R^2 = 0.50$ instead of 0.18). What does the 1.95 elasticity of nominal exports with respect to the real effective exchange rate mean in real terms? Referring to the above discussion on the degree of exchange-rate pass-through, the unobserved index of export prices in foreign currency units shrinks between 0 and 1 percent in case of a one percent real depreciation of the hryvnia.[20] The higher profitability of exporting activities relative to domestic sales thus triggers a long-run response in real export supply of between 1.95 and 2.95 percent.[21] The results are therefore in line with an upward-sloping long-run supply curve. However, as the slope of the latter also seems to depend on the specification of the dynamic adjustment[22], which is not discussed here due to the lacking distinction between prices and quantities, it is difficult to say whether 2.95 is high or low by international standards. So the only distinction made is between flat and upward-sloping.

A definite strength of our results is that both in section 6.3 and in 6.4 almost precisely the same long-run demand vector is found and that this vector corresponds to the one in the previous section (Table 3). Thus the shape of demand

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[20] Only if the depreciation stems from foreign prices rising faster than domestic ones would we also expect an increase in the "dollar" export price by up to one percent holding Ukraine's competitive position constant. But this case is not the usual one in the period 1993–2000.

[21] Relative profitability of exports is seen as the key determinant of export supply in the conventional literature (Goldstein and Khan 1978: 276 and 1985: 1060–61, Sawyer and Sprinkle 1999: 10).

[22] Browne (1982: 346–47) argues that long-run supply should be relatively steep for small open economies because a high export share in GDP means that even strong world price incentives only yield a relatively modest percentage growth in production of exportables due to resource constraints. Unlike Goldstein and Khan (1978) Browne uses a dynamic specification where export prices adjust to demand imbalances while quantities converge to suppliers' desired levels, and finds a much steeper supply curve for Ireland.

for nominal Ukrainian exports is robust with respect to a wide array of thinkable model specifications. As to the upward-sloping supply curve, the trend coefficient is estimated to be 0.01 suggesting that exporters increased export prices in foreign currency units by roughly 4 percent per year during the estimation period to cope with the severe and permanent real appreciation of the national currency (amounting to an average 9.6 percent per year, cf. Table 6, line 5). As discussed earlier, this intermediate strategy might have been a loss-minimizing device between the plagues of declining competitiveness and the one of falling (or even negative) profit margins. Although these results fit well into economic reasoning, some open questions remain, especially with respect to the loading coefficients. For instance, nominal exports themselves have no role to play in the adjustment of export demand or price imbalances, the corresponding loadings are insignificant. One then recurs to the significant loadings in equation [3]: too low exports lead to a real depreciation unlike in the previous section.[23] The positive coefficient on the second cointegration vector suggests that too high an export price level leads to devaluation. Yet absolute values exceeding unity for both loadings in [3] point to stability problems in the supply 1 specification.[24]

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[23] However, see footnote 13. The finding in Table 9 seems in line with the literature on the effects of current account imbalances on the exchange rate. In case of sustained current account deficits market participants may get convinced that the “equilibrium” exchange rate (the one restoring a sustainable net foreign asset position of the country) has risen (Hooper and Morton 1982). Another theory suggests that high external debt of a country may lead risk-averse foreign investors to demand a risk-premium for holding the country’s assets such that uncovered interest parity no longer holds (Adler and Dumas 1983).

[24] As to equation [2], the loadings are well-signed although implausibly high (both around 0.40). Excessive exports stimulate production and income abroad and too high export prices put a strain on foreign businesses.

Table 9: Estimation results from the Ukrainian export system with Supply 1[a,b]

Vector 1 ("Export demand"):

$$s = y^{*} + \underset{(5.58)}{1.73}e$$

Test for over-identifying restrictions: $c^2(1) = 1.68[0.20]$

Vector 2 ("Supply 1":"Unobserved export prices"):

$$s = \underset{(6.87)}{1.95}e + \underset{(10.00)}{0.01}t$$

$$\Delta s_t = \underset{(0.56)}{0.59} + \underset{(0.84)}{0.29}\left[s_{t-1} - \left(y^{*}_{t-1} + 1.73e_{t-1}\right)\right] - \underset{(-1.07)}{0.40}\left[s_{t-1} - \left(1.95e_{t-1} + 0.01t\right)\right]$$

[1] $$- \underset{(-2.75)}{0.37}\Delta s_{t-1} + \hat{u}_{1t}$$

$$(\text{STDDEV} = 0.061;\ R^2 = 0.57;\ \text{normality} = 0.75;\ \text{Arch}(2) = 0.91)$$

[2] $$\Delta y^{*}_t = \underset{(4.08)}{1.37} + \underset{(3.74)}{0.41}\left[s_{t-1} - \left(y^{*}_{t-1} + 1.73e_{t-1}\right)\right] - \underset{(-3.51)}{0.42}\left[s_{t-1} - \left(1.95e_{t-1} + 0.01t\right)\right] + \hat{u}_{2t}$$

$$(\text{STDDEV} = 0.020;\ R^2 = 0.70;\ \text{normality} = 1.79;\ \text{Arch}(2) = 1.12)$$

$$\Delta e_t = \underset{(-4.16)}{-6.17} - \underset{(-4.30)}{2.07}\left[s_{t-1} - \left(y^{*}_{t-1} + 1.73e_{t-1}\right)\right] + \underset{(4.40)}{2.31}\left[s_{t-1} - \left(1.95e_{t-1} + 0.01t\right)\right]$$

[3] $$- \underset{(-1.82)}{1.35}\Delta y^{*}_{t-1} + \hat{u}_{3t}$$

$$(\text{STDDEV} = 0.086;\ R^2 = 0.50;\ \text{normality} = 1.96;\ \text{Arch}(2) = 1.44)$$

Multivariate statistics[b]:

$L - B(8) = 56.21[0.19],\ LM(1) = 13.26[0.15],\ LM(4) = 7.74[0.56],$

normality : $c^2(6) = 4.15[0.66]$

[a] Export prices do not depend on foreign production in the long run; *k=2*; *r=2*; unrestricted constant plus trend restricted to the cointegration space. — [b] All footnotes of Table 3 also apply to this table.

6.4. Final Specification With An Infinitely Elastic Long-Run Supply Curve

In this section I stick to the infinitely-elastic-supply hypothesis widely accepted in the empirical literature (Goldstein and Khan 1985: 1087; Sawyer and Sprinkle 1999: 10) to see if the specification problems just discussed get settled. They do as one can see in Table 10. The alternative specification is implemented by restricting the influence of nominal exports to zero in the long-run.[25] Normalizing on e shows that the second vector can be read as a long-run real effective exchange rate relationship given elastic export supply. The influence of foreign production is positive as suggested by the Harrod-Balassa-Samuelson effect (Harrod 1933; Balassa (1964); Samuelson (1964)): an increase in foreign production is likely to imply a positive growth differential between foreign countries and Ukraine, which ultimately leads to a real appreciation of foreign currencies relative to the hryvnia.[26] Nevertheless it was Ukraine which suffered substantial real appreciation due to high domestic inflation and other transformation-related specific problems. The latter are captured by the positive coefficient of the time trend.

How do this real exchange rate vector and the export demand vector interact in the dynamic system? Again the focus is on the loading coefficients as most of the short-term coefficients are insignificant. Nominal exports are found to react only to demand imbalances; the second vector has no significant impact on exports. Excess exports (indicated by a positive error correction term in the demand relationship) trigger a downward correction of exports themselves and a less intuitive rise in the real effective exchange rate, as in section 4. If the excess is caused by export volumes, the mechanism might run via a foreign production stimulus resulting from excessive Ukrainian deliveries; the positive reaction in

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[25] Strictly speaking, when only nominal exports are known, one cannot distinguish between a horizontal and a vertical supply curve without an assumption on the long-run export price reaction to a change in the real effective hryvnia exchange rate because $\lim\limits_{p_x \longrightarrow \infty} p_x x = \lim\limits_{x \longrightarrow \infty} p_x x \longrightarrow \infty$. We rule out the possibility of a vertical supply curve *assuming* that export prices do not rise or fall infinitely in case of a change in e. This is equivalent to saying that investment bottlenecks do not hamper investment forever (when e rises) and that market exit of uncompetitive firms is possible in the long run (when e falls).

[26] The point is succinctly developed in Obstfeld and Rogoff (1996: 210–213).

*y** resulting from too high a real exchange rate level (second loading coefficient in [2]) points into a similar direction suggesting that Ukraine's exports serve as inputs into foreign production and thus are rather complements than substitutes to

Table 10: Estimation results from the Ukrainian export system with Supply 2[a,b]

Vector 1 ("Export demand"):

$$s = y^{*} + \underset{(6.24)}{1.75}e$$

Test for over-identifying restrictions: $c^{2}(1) = 1.68[0.20]$

Vector 2 ("Supply 2", "Real exchange rate given horizontal export supply"):

$$e = \underset{(14.85)}{4.54}y^{\prime *} + \underset{(-10.75)}{0.043}t$$

[1]

$$\Delta s_t = \underset{(0.56)}{0.59} - \underset{(-3.08)}{0.11}\left[s_{t-1} - \left(y^{*}_{t-1} + 1.75e_{t-1}\right)\right] + \underset{(1.07)}{0.09}\left[e_{t-1} - \left(4.54y^{*}_{t-1} + 0.043t\right)\right] - \underset{(-2.75)}{0.37}\Delta s_{t-1} + \hat{u}_{1t}$$

$(\text{STDDEV} = 0.061;\ R^2 = 0.57;\ \text{normality} = 0.75;\ \text{Arch}(2) = 0.91)$

[2]

$$\Delta y^{*}_t = \underset{(4.09)}{1.38} - \underset{(-0.89)}{0.01}\left[s_{t-1} - \left(y^{*}_{t-1} + 1.75e_{t-1}\right)\right] + \underset{(3.51)}{0.09}\left[e_{t-1} - \left(4.54y^{*}_{t-1} + 0.043t\right)\right] + \hat{u}_{2t}$$

$(\text{STDDEV} = 0.020;\ R^2 = 0.70;\ \text{normality} = 1.79;\ \text{Arch}(2) = 1.12)$

[3]

$$\Delta e_t = \underset{(-4.16)}{-6.17} + \underset{(4.72)}{0.24}\left[s_{t-1} - \left(y^{*}_{t-1} + 1.75e_{t-1}\right)\right] - \underset{(-4.40)}{0.51}\left[e_{t-1} - \left(4.54y^{*}_{t-1} + 0.043t\right)\right] - \underset{(-1.82)}{1.35}\Delta y^{*}_{t-1} + \hat{u}_{3t}$$

$(\text{STDDEV} = 0.086;\ R^2 = 0.50;\ \text{normality} = 1.96;\ \text{Arch}(2) = 1.44)$

Multivariate statistics[b]:

$L - B(8) = 56.20[0.19],\ LM(1) = 13.29[0.15],\ LM(4) = 7.71[0.56],$
normality : $c^{2}(6) = 4.15[0.66]$

[a] The real exchange rate does not depend on exports in the long run (horizontal export supply curve); k = 2; r = 2; unrestricted constant plus trend restricted to the cointegration space. — [b] All footnotes of Table 3 also apply to this table.

foreign goods and factors. But as overshooting in case of nominal exports may also stem from export price increases, the positive first loading coefficient in equation [3] may as well suggest export price hikes lead to a depreciation in the long run. In any case, an upward spiral in both *s* and *e* seems to be prevented by a large value of the second loading in [3] (–0.51) which brings *e* back towards where it should be according to the second vector. Without being exogenous to the system, foreign industrial production at least seems to be exogenous with respect to the export demand relationship, which is a plausible result for a relatively small economy as Ukraine.

7. Summary and Conclusions

In this paper the determinants of Ukrainian exports are estimated by means of cointegration analysis. In the standard specification, which treats the (infinitely elastic) long-run supply by assumption, one cointegration vector in line with standard production theory is found. Due to non availability of export prices one has to make assumptions as to the long-run reaction of export prices to changes in the real exchange rate. Accounting for all plausible long-run price reactions, the cautious conclusion is that real demand for Ukrainian exports is price-elastic with the estimated long-run coefficients being a lower bound of the economically relevant price-elasticity of real demand.

High and persistent domestic inflation especially at the beginning of the estimation period also justify the rather unconventional specification which allows for a deterministic trend in the cointegration space. The reasons are that Ukrainian cost-push inflation has arguably brought about a systematic rise in unobservable export prices well above the one observed in foreign producer prices, on the one hand and a substantial and permanent real appreciation of the hryvnia, on the other hand, which are not captured elsewhere in the regressors. The richer specification leads to detection of a second vector which is best identified as a structural real exchange rate relationship with infinitely elastic export supply. While the interpretation of the second vector is somewhat ambiguous (horizontal versus upward-sloping export supply), the same demand vector is found to hold in all specifications and lends the analysis an astonishing degree of robustness.

The causes of Ukraine's poor export performance during the nineties are by now clear. The proximity of trading partners suffering from similar problems of economic transition and slightly shrinking aggregate production, the demand for Ukrainian exports would — at best — have stagnated in an environment of stable terms of trade. The policy implication is straightforward: If the country modifies the geographical orientation of trade strengthening its trade links with fast-growing economies at the cost of those with states of the former Soviet Union, it will become less dependent on sluggish growth in the Eastern neighborhood. WTO accession could be a trigger pushing the country towards integration into the international division of labor dictated by principles of comparative advantage. This would probably lead to tough structural change with winning and losing industries but aggregate exports are very likely to benefit (European Commission 2001a).[27]

The second brake on Ukrainian exports have been hyperinflation and the resulting real appreciation of the hryvnia at the beginning of the observation period. As goods prices generally rise rapidly to at least partly converge to international

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[27] In case of no WTO accession the country would largely benefit from a preferential trade agreement with the EU as two-thirds of total exports are sensitive according to EU definitions (Shpek 2000: 22)

prices (Cushman et al. 2001: 254–255) higher inflation in Ukraine compared to Central and Western Europe seems unavoidable for the future, as well. But what is needed is a maximum stability in the conduct of monetary policy which would allow for a controlled nominal depreciation keeping pace with the international inflation differential and thus holding the real effective exchange rate of the hryvnia approximately constant. Given the high price-sensitivity of the Ukrainian export assortment, a stable monetary framework accompanied by productivity increases would promise the fastest improvements in export growth.

Appendix: Unit root tests to determine the order of integration

The results of the unit root tests according to Dickey and Fuller (1981) are summarized in Table A1. All test equations contain an intercept and a linear time trend except the real effective hryvnia exchange rate. For the latter a long-run deterministic trend is not necessarily plausible both on economic grounds and upon visual inspection of the data. Therefore both the results for the model with trend and intercept and for the model with an intercept alone are reported. The number of lags in the test equations is chosen minimal subject to the requirement of freedom of autocorrelation up to the fourth order.

The results show that all time series used are non stationary. Nominal exports and the real effective exchange rate clearly turn out to be integrated of order 1, while mixed results are obtained for the index of foreign industrial production. Here the ADF-test on first differences might be of little reliability because of the high number of lagged second differences needed to obtain white noise in the residuals. Therefore the result of the Phillips-Perron (1988) unit root test is added which is robust against autocorrelation and heteroskedasticity in the residuals. As the null of y^* being I(2) is rejected, I cautiously conclude that all time series are integrated of order one.

As it is empirically very difficult to distinguish between deterministic and stochastic trends in small samples (Harris 1995: 39), KPSS-tests for stationarity[28] of levels and first differences are run as a robustness check of the results. Unlike the ADF and Phillips-Perron tests, the null hypothesis is stationarity of the series under investigation. Nevertheless in the case at hand the KPSS-test leads to the same conclusions as the ADF-test.[29]

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[28] See Kwiatkowski, Phillips, Schmidt and Shin (1992) and Shin (1994) for methodology and critical values.

[29] The results are available from the authors upon request.

Table A.1: Results of the Augmented Dickey Fuller (ADF) Unit-Root Tests

Variable	Test for I (0)[a]		Test for I (1)[a]		Result
	Specification[b]	ADF test statistics[c]	Specification[b]	ADF test statistics[c]	
s	T, 0	–2.23	C, 0	–5.43***	I (1)
*y**	T, 3	–2.77	C, 7	–1.81	I (2)[d]
e	T, 0	–1.56	C, 0	–5.91***	
	C, 0	–2.38	N, 0	–5.72***	I (1)

[a]***(**,*)means rejection at the 1 pc (5 pc, 10 pc) significance level. — [b]T: model with drift and trend; C: model with drift; N: model without drift and trend. The figure indicates the number of lagged variables in the test equation. — [c]Augmented Dickey-Fuller *t*-test. — [d]The Phillips-Perron test (with truncation lag 3) yields a test statistic of –2.68, which allows for rejection of the null of non stationarity at the 10 percent level (critical value: –2.62).

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