The European Journal of Comparative Economics
Vol. 1, n. 2, 2004, pp. 175-193
ISSN 1824-2979

Price and Income Elasticities of Russian Exports

Bernardina Algieri[1]
Università della Calabria & Universität Bonn

Abstract

The paper gauges export demand elasticities for Russia using an Error Correction technique within a cointegration framework. An extended version of the Imperfect Substitutes Model has been implemented to estimate the sensitivity of Russian exports without oil components to price and to Russian and world income. Our results suggest a robust and negative long run cointegration relationship between the real effective exchange rate, defined as the weighted average of the rouble's exchange rates versus a basket of the three currencies with the largest share in the trade turnover adjusted to incorporate inflation rate differences (the ratio of the domestic price indices to the foreign price indices), and Russian exports. An increase in exports by 24% is caused by a real depreciation by 10%. Furthermore, a 10% growth in world income leads to a 33% rise in exports. Finally, exports drop by 14% whenever a 10% increase in domestic income occurs.

JEL Classification: C22, F19, P27

Keywords: Russia, export demand function, elasticities, cointegration.

1. Introduction

The aim of this paper is to evaluate the role played by income and prices in the determination of Russian exports. In particular, it will calculate to what extent changes in prices affect Russian exports and to what extent changes in foreign and Russian income impact on the demand for Russian products. Income and price elasticities are estimated within a Cointegration framework using the Error Correction Model (ECM) technique. It is important to estimate price and income elasticities because they can be applied to many relevant macro-economic policy issues: the effect of both monetary and fiscal policies and expenditure switching policies (such as exchange rate, subsidy and tariff policies) on a country's balance of payments, the impact of external balance restrictions on domestic policy measures, the international transmission of changes in economic activity and prices and the employment effects of changes in own or partner-countries' trade restraints.

Substantial empirical literature exists on the estimation of price and income elasticities in international trade, much of it focused on U.S and European trade. Most econometric estimations indicate that price elasticities fall in a range of 0 to –4.0, while income elasticities fall between 0.17 and 4.5. Since the values of price elasticities vary considerably, the recent literature questions the effectiveness of real devaluation in affecting exports and imports. According to Rose (1990, 1991) and Ostry and Rose (1992), a real depreciation does not impact significantly on the trade balance. Reinhart (1995), Senhadji and Montenegro (1998), Senhadji and Montenegro (1999) provide instead, strong support to the view that depreciations improve the trade balance. It seems that low econometric estimates of price elasticities are unreliable for the purpose

[1] Corresponding Author: Bernardina Algieri, Via della Resistenza 70, 87040 Castrolibero, Cosenza, Italy. E-mail: b.algieri@unical.it. I wish to thank Prof. Antonio Aquino, University of Calabria for extensive discussions and suggestions. I am grateful to Dr. PD Peter Wehrheim, University of Bonn, to Dominic Mulreany, to Prof. Michael Keren and to an anonymous referee for their helpful comments.

of forecasting the effect of a depreciation, and there is a strong presumption that these elasticities lead to a considerable underestimation of its effectiveness.

The paper is divided into seven sections. Section 2 briefly reviews the trade modelling literature. Section 3 describes the imperfect substitute model. Section 4 specifies a model for Russia and provides an explanation of data employed. Section 5 shows the empirical analysis within a Cointegration framework. The main findings are presented in section 6. Section 7 concludes.

2. Trade Modelling

The behaviour of foreign trade flows has been subjected to many empirical investigations through the estimation of trade equations. The latter are equations for the time-series behaviour of the quantities and prices of imported and exported goods. Early estimations of income and price elasticities have been investigated and assessed by Prais (1962). Early world trade models are examined in Taplin (1973). Multi-country models have been gauged by Deardorff and Stern (1978). Special attention should be paid to trade surveys by Leamer and Stern (1970), Stern et al. (1976) and to the works by Chipman (1985), Goldstein and Khan (1985), Faini et al. (1992), Hung et al. (1993), de la Croix and Urbain (1995), Hooper et al. (1998), Marquez and McNeilly (1998), Senhadji and Montenegro (1999), Nielsen (2001), Banco de Espaňa (2003).

The question of how the time series behaviour of imports and exports should be modelled has been subject to much debate. The appropriate model relies on: the type of traded commodity, i.e. if it is a homogeneous or a differentiated good; the main purpose to which the traded product is destined, i.e. if it used as factor of production or as final good; the institutional and legal structure under which trade takes place; the aim of the modelling analysis, i.e. if it is necessary to forecast or to test hypotheses; and the availability of data, i.e. if data are annual or quarterly or if they are disaggregated or aggregated.

The empirical literature has been characterised by two general models of trade, namely, the imperfect substitutes model and the perfect substitutes model. The two models have often been considered as competitors, because most trade analyses have gauged aggregate exports and/or imports. When aggregation is no more a severe constraint and it is possible to disaggregate data, the two aforementioned models could be viewed as complements: "one concerning trade for differentiated commodities and the other regarding trade for close, if not perfect, substitutes" (Goldestein and Khan, 1985; Senhadji and Montenegro, 1998).

3. The Economic Model: The Imperfect Substitutes Model

Since the amount of export-import adjustments depends on the sensitivity to price and income variations, it is relevant to calculate the price and income elasticity of a country's export volumes. The theoretical foundation of the empirical analysis is the Imperfect Substitutes Model. The basic assumption of the model is that neither imports nor exports are perfect substitutes for domestic products. Such a hypothesis is confirmed by empirical evidence. If domestic and foreign goods were perfect substitutes, a given country would be either an exporter or an importer. Since the world market is characterised by the presence of bilateral trade and the coexistence between imports and domestic production, the hypothesis of perfect substitution can be rejected. Moreover, a large body of empirical studies (Lipsey (1978); Kravis and Lipsey (1983);

Giovannini (1988); Wolf et al. (1994, 2000),) have shown that price differentials can be surprisingly large for the same product in different countries, as well as between the domestic and export prices of a given product in the same country. In other words, the "law of one price" fails dramatically in practice, even for products that commonly enter international trade. It seems therefore, that finite price elasticities of demand and supply (as the imperfect substitutes models postulates) can in fact be estimated for most traded goods.

The imperfect substitutes model (Goldstein, Moris, Khan 1985; Marquez and McNeilly, 1988; Hooper and Marquez, 1995) of the home country's exports to, and imports from, the rest of the world (*) is formalized by a set of equations:

$M^d = \gamma\ (Y, P^M, P,)$	$\gamma_1, \gamma_3 > 0,\ \gamma_2 < 0$	(1)
$X^d = \pi(Y^*\ e, P^X, P^*e)$	$\pi_1, \pi_3 > 0, \pi_2 < 0$	(2)
$M^s = \varphi(\ P^{M^*}\ (1+S^*), P^*)$	$\varphi_1 > 0, \varphi_2 < 0$	(3)
$X^s = \xi(P^X\ (1+S), P)$	$\xi_1 > 0, \xi_2 < 0$	(4)
$P^M = P^{X^*}\ (1+T)e$		(5)
$P^{M^*} = P^X\ (1+T^*)/e$		(6)
$M^d = M^s e$		(7)
$X^d = X^s$		(8)

The eight equations identify the quantities of imports demanded by the home country (M^d), the quantity of exports demanded by the world from the home country (X^d), the quantity of imports supplied by the rest of the world to the home country (M^s), the quantity of the home country exports to the rest of the world (X^s), the prices in domestic currency paid by the importers (P^M and P^{M*}) and the prices in domestic currency paid to the exporters (P^X and P^{X*}). The level of nominal income (Y, Y*), the prices of domestic commodities produced within the regions (P, P*), proportional tariffs (T, T*), subsides to imports and exports (S, S*) and the real exchange rate (e) are the explanatory variables.

The main features of the imperfect substitutes model can be summed up as follows. Along with the standard demand theory, it is supposed that the representative agent maximises his lifetime utility subject to a lifetime budget constraint. The resulting demand functions for exports and imports therefore, describe the quantity demanded as a function of the level of monetary income in the importing country, the imported product's own price, and the price of domestic substitutes. By considering a logarithmic utility function, the income (γ_1 and π_1) and price elasticity (γ_3 and π_3) of substitutes are assumed to be positive, while the price elasticity of the traded product is assumed to be negative (γ_2 and π_2).

Let us assume the demand function to be homogeneous of degree 0, equation 1 can be written in the following way:

$$M^d = \gamma\ (Y/P\ , P^M/\ P) \qquad \gamma'_1 > 0,\ \gamma'_2 < 0$$

Where Y/P = real income
$P^M/\ P$ = real import price

Considering an n-country model, the symmetry between the demand function for imports and the demand function for exports vanishes. Imports compete, in fact, only with goods produced within the country. Exports compete both with goods produced in the imported country and with exported goods by third countries. The equation 2 is corrected with prices of competing goods.

$X^d / X_*^d = \pi (P^* X^d / P^* X_*^d)$

where X_*^d is the demand for exports to the rest of the world from third countries.

The supply functions depend on the prices of exported and domestic goods and on subsidies. The price elasticities of exported and local commodities (φ_1 and ξ_1) are assumed to be positive, the price elasticities of substitutes (φ_2 and ξ_2) are supposed to be negative. The equilibrium conditions are represented by the last two equations. The implicit hypothesis is that prices move in order to equate demand and supply over time.

The imperfect substitutes model, by presenting both demand and supply side equations, allows to identify simultaneous relationships among quantities and prices. In spite of this peculiar characteristic, brilliantly pointed out by Orcutt, (1950) and Goldstein and Kahn (1985), a host of time series works on export and import equations have considered the supply side only by assumption. Throughout the early 90s, the standard methodology to estimate import and export demand (eq. (1) and eq. (2)) was based on the assumption of an infinite supply-price elasticity for imports and exports (φ_1 in eq. (3) and ξ_1 in eq. (4)). Under this hypothesis, P^M and P^X were viewed as exogenous and thus estimated by single equations. If the supply elasticities were instead, less than infinite, the problem would be more cumbersome because one should either appraise the complete structural system of simultaneous equations or solve the reduced form for quantities and prices as functions of the exogenous variables in the system. After 1995, economic researchers have used either cointegration analyses to cope with simultaneity problems or fully-modified-OLS methodologies to overcome endogeneity and serial correlation biases.

4 A Model for Russia

Let us suppose that Russia (the exporting country) has only one trading partner (the rest of the world). Russia's export demand (x_t) will hence coincide with the import demand of the rest of the world (m_t*). We continue to assume that there is a representative agent in the rest of the world, who lives forever and maximises his utility by choosing how much to consume of his domestic endowment (w_t*) and of the imported good (m_t*). It is supposed further, that there is no production sector, because production often involves the combining of intermediate inputs by using factors of production, while the model makes no distinction between intermediate and final products. The representative consumer in the rest of the world maximises an inter-temporal utility function over time (Ut) expressed as:

$$\underset{\left\{w_t^*, m_t^*\right\}_{t=0}^{\infty}}{Max} U_t = \underset{\left\{w_t^*, m_t^*\right\}_{t=0}^{\infty}}{Max} E_t \cdot \left\{ \sum_{t=0}^{\infty} (1+\delta)^{-1} \cdot u(w_t^*, m_t^*) \right\} \qquad (9)$$

subject to his budget constraint:

$$b_{t+1}^* = (1+r) \cdot b_t^* + (s_t^* - w_t^*) - p_t \cdot m_t^* \qquad (10)$$

and to a transversality condition, which excludes Ponzi-games, i.e. the fact that a consumer can freely consume all lifetime resources, by borrowing forever without extinguishing his debt.

$$\lim_{T \to \infty} \frac{b^*_{T+1}}{\prod_{t=0}^{T} (1+r)^{-1}} = 0 \qquad (11)$$

All the starred variables denote the rest of the world (the importing country) while non-starred variables refer to Russia. E{·} is the expectation operator at time t; δ is the consumer's rate of time preferences, i.e. the subjective discount rate, which measures the individual's impatience to consume; agents are free to borrow and lend at the same world interest rate *r,* that is the yield on capital; b^*_{t+1} denotes the next period stock of Russian bonds held by the rest of the world if positive and the next period stock of foreign bonds held by Russia if negative; p_t is the price of the Russian good in terms of foreign commodity; s^*_t is the stochastic endowment which follows an AR(1) process of the form:

$$s^*_t = \mu \cdot s^*_{t-1} + (1-\mu) \cdot \bar{s}^* + \varepsilon^*_t$$

$$0 \le \mu \le 1 \qquad (12)$$

$$\varepsilon^*_t \approx \left(0, \sigma^2\right)$$

with an unconditional mean $\bar{s}^*$ and an unconditional variance $\sigma^2/(1-\mu^2)$. ε^*_t is an independent and identically distributed shock to the stochastic endowment with zero mean and variance σ^2 . μ governs the degree of persistence of the endowment shock. The first order conditions for the individual's problem are:

$$\frac{\partial L_t}{\partial w^*_t} : \; u_{w^*}(t) - \lambda_t = 0 \qquad (13)$$

$$\frac{\partial L_t}{\partial m^*_t} : \; u_{m^*}(t) - \lambda_t \cdot p_t = 0 \qquad (14)$$

$$\lambda_t = (1+\delta)^{-1} \cdot E_t (1+r) \cdot \lambda_{t+1} \qquad (15)$$

where L is the Lagrangian and λ_t is the Lagrange multiplier associated with the budget constraint. Consider the case in which the individual utility function is of addilog type, as in Ogaki (1992) and Clarida (1994), de la Croix and Urban (1995), Senhadji and Montenegro (1999). The specific form of the instantaneous utility function is:

$$u_t(w^*_t, m^*_t) = \frac{\Omega_t \cdot (w^*_t)^{1-\alpha}}{1-\alpha} + \frac{\Psi_t \cdot (m^*_t)^{1-\beta}}{1-\beta} \qquad (16)$$

$$\Omega_t = e^{a_0 + \upsilon_{\Omega,t}} \qquad (17)$$

$$\Psi_t = e^{b_0 + \upsilon_{\Psi,t}} \qquad (18)$$

where Ω_t and Ψ_t are exponential stationary random shocks, which cause variations in the preferences of the representative agent, $\upsilon_{\Omega,t}$ and $\upsilon_{\Psi,t}$ are stationary shocks, α and β are called curvature parameters and their inverse can be interpreted as long-run intertemporal elasticities of substitution between the domestic and the imported good. Solving the maximization problem for the explicit utility function (16) and substituting the values for Ω_t and Ψ_t we will have the following first order conditions:

$$w_t^* = \lambda_t^{-\frac{1}{\alpha}} \cdot (e^{a_0+\upsilon_{\Omega,t}})^{\frac{1}{\alpha}} \quad (19)$$

$$m_t^* = \lambda_t^{-\frac{1}{\beta}} \cdot (e^{b_0+\upsilon_{\Psi,t}})^{\frac{1}{\beta}} \cdot p_t^{-\frac{1}{\beta}} \quad (20)$$

Taking the log for the equations (19) and (20), solving the equation (19) for λ and substituting it in the equation (20) yields:

$$log\ m_t^* = \frac{\alpha}{\beta} log\ w_t - \frac{1}{\beta} log\ p_t - \frac{a_0}{\beta} - \frac{\upsilon_{\Omega,t}}{\beta} + \frac{b_0}{\beta} + \frac{\upsilon_{\Psi,t}}{\beta} \quad (21)$$

posing $c_0 = \frac{1}{\beta}(b_0 - a_0)$ and $\upsilon_t = \frac{1}{\beta}(\upsilon_{\Psi,t} - \upsilon_{\Omega t})$ leads to the standard export demand function of the economic imperfect substitutes model:

$$log\ m_t^* = log\ x_t^* = c_0 + \frac{\alpha}{\beta} log\ w_t - \frac{1}{\beta} log\ p_t + \upsilon_t \quad (22)$$

4.1 The extended Imperfect Substitutes Model

To take into account the availability of output as a determinant of exports, Russian national income (y_t) has been added to the explanatory variables of the export demand function considered in the imperfect substitutes model. The expected sign of the coefficient of the aforementioned variable is however ambiguous, since the value of national income stems from the equality between production and effective demand. If increases in national income are mainly caused by upsurges in domestic Russian demand, exports drop because of a slump in the availability of goods for exports. Conversely, if increases in national income are mainly pushed by a surge in Russian production capacity, exports could rise due to a greater availability of commodities (Aquino, 1982; Kalecki, 1971).

We have therefore, estimated the coefficients of the extended "stochastic" econometric specification.

$$log\ m_t^* = log\ x_t^* = c_0 - \frac{1}{\beta} log\ p_t + \frac{\alpha}{\beta} log\ w_t + \zeta\ log\ y_t + \upsilon_t \quad (23)$$

where υ_t is an iid disturbance:
$\upsilon_t \sim N(0, \sigma^2)$

4.2 The Data

The quantitative stochastic equation for Russia has been constructed, using "Russian Economic Trends", CBR data and the "Indicators of Industry and Services" by OECD. In particular, the real effective exchange rate with base 1995=100 has been used for variable p_t. The real effective exchange rate is a trade weighted exchange rate, whose weightings are 40% the US, 40% Germany and 20% Ukraine. Thus, the real effective exchange rate is a significant indicator which measures the variations in competitiveness of the Russian economy. An increase in this series represents a real appreciation. y_t is the Russian real income at 1995 prices. In order to avoid endogeneity and sign problems, y_t does not include Russian exports. w_t is the OECD total industries production with base 1990=100. x_t are the Russian exports values with base 1995=100. To avoid aggregation problems, we have excluded the exports of oil and its products from the total exports values[2]. Monthly data ranging from December 1993 to November 2001 have been used. The reasons to adopt monthly data are mainly two: 1) a technical reason concerns the shortage of years passed since the USSR dissolution, hence we have a lack of sufficient observations regarding quarterly data 2) a non technical reason regarding the rapid changes which occur in the Russian economy and as a consequence there is an interest in monitoring them.

5 Empirical Analysis: Econometric Characteristics of the Time Series

The first critical step of the analysis is to visually inspect the data. The variables' dynamics are sketched in Appendix Figure 2. They have been seasonally adjusted to account for their seasonal movements.

Each series seems to meander in a fashion characteristic of a random walk. To formally test for the presence of a unit root in the export–elasticity function, the augmented Dickey Fuller test (ADF) and the Phillips-Perron test (PP) have been implemented for each variable. The lag structure in the ADF regression is selected using the Akaike Information Criterion (AIC). The model includes a trend, since its presence is clear from the previous graphical inspection. The results are reported in Appendix

Table 4 .

The variables are not stationary, in fact Ho (existence of unit root) cannot be rejected. The critical values for the rejection of the hypothesis of a unit root are those computed according to the MacKinnon criterion. The variables are integrated of order one I(1) at 1%, 5% and 10% critical values, with the exception of *logx* which is stationary according to the Phillips-Perron test. Following the augmented D-F procedure all the variables are I(1) at each critical level.

5.1 The Pesaran-Shin-Smith Approach

Since available time series are relatively short, the rather new approach by Pesaran, Shin and Smith (2001) has been adopted, in the attempt to obtain robust estimates. In fact, as an anonymous referee pointed out and as Campbell and Perron (1991, pag.13) put forward:

2 In a first analysis we estimated aggregate exports. Clearly, oil dominated the series, i.e. although oil price fluctuated quite wildly, the measured elasticity was very low because oil exports were not price sensitive.

> "It turns out that for tests of the unit root hypothesis versus stationary alternatives the power depends very little on the number of observations per se but is rather influenced in an important way by the span of the data".

This objection renders the conducted unit root testing less reliable and makes the Pesaran, Shin and Smith approach (PSS) desirable.

The PSS is a bound test for analysing the following long run relationship:

$$\begin{aligned} \text{dlog x} &= \alpha + \beta_1 \log x_{t-1} + \beta_2 \log p_{t-1} + \beta_3 \log y_{t-1} + \beta_4 \log w_{t-1} + \sum_{i=1}^{4} \chi_i d \log x_{t-i} + \\ &+ \sum_{i=1}^{4} \eta_i d \log p_{t-i} + \sum_{i=1}^{4} \varphi_i d \log y_{t-i} + \sum_{i=1}^{4} \gamma_i d \log w_{t-i} + \upsilon_t \end{aligned} \tag{24}$$

The equation contains *difference in lags* (d) of the dependent and independent variables and *lags* one period of explanatory variables (eq. 24). Each *difference in lags* of the dependent and the independent variables constitutes the short run dynamics. They describe how the dependent variable is changed by the first difference of its own lagged values and the first differences of the lagged independent variable. Each *lag* and namely, the ratios $\beta_2/(-\beta_1)$, $\beta_3/(-\beta_1)$ $\beta_4/(-\beta_1)$ constitute the long run dynamics.

The PPS Autoregressive Distributed Lag (ARDL) method has the advantage of avoiding the classification of variables into *I*(1) or *I*(0) and unlike standard cointegration tests, there is no need for unit root pre-testing. Two stages are involved in the PSS procedure. First, the null hypothesis of no cointegration is tested against the alternative of cointegration:

Ho: $\beta_1=\beta_2=\beta_3=\beta_4=0$ vs. H1: $\beta_1\neq\beta_2\neq\beta_3\neq\beta_4=0$

by using the *F*-test or Wald test. Since the asymptotic distribution of this *F*-statistic is non-standard, Pesaran et al. (2001) have tabulated two sets of appropriate critical values. One set assumes all variables are *I*(1) and another assumes that they are all *I*(0). This provides a band covering all possible classifications of the variables into *I*(1) and *I*(0) or even fractionally integrated. If the calculated *F*-statistic lies above the upper level of the band, the null is rejected, indicating cointegration. If the calculated F statistic falls below the lower level of the band, the null cannot be rejected, supporting lack of cointegration. If, however, it falls within the band, the result is inconclusive. If there is evidence of co-integration, a second stage involves defining the error correction term. This is undertaken in an analogous way to the Eagle and Granger procedure (paragraph 5.2).

We carried out the first stage by imposing four lags (i=4) on each first difference term in the ARDL model. The appropriate length of the distributed lags was chosen by adopting the Akaike Information Criterion and the Schwarz Criterion.

An *F*-statistic of 6.04 was obtained when a trend term was included in the model. This is greater than the upper level of the critical band (i.e., 5.07) supporting cointegration. The PSS estimates are reported in Table 1 .

Table 1 Pesaran Shin Smith Test

Null Hypothesis:	$\beta_1=0$ $\beta_2=0$ $\beta_3=0$ $\beta_4=0$		
F-statistic	6.047484	Probability	0.000272
Chi-square	24.18994	Probability	0.000073

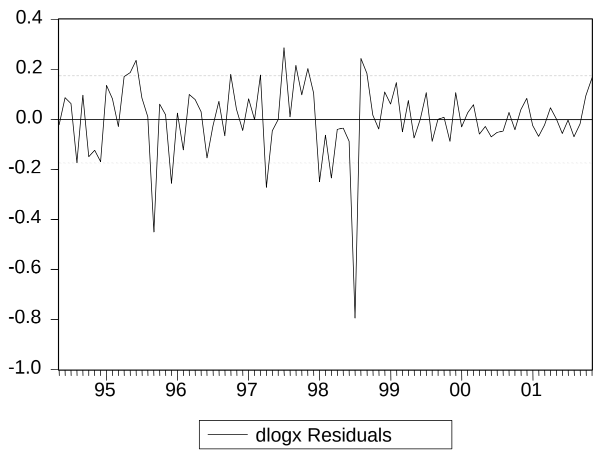
Pesaran critical values 5% significance
I(0)=4.01 I(1)=5.07

5.2 Cointegration Analysis

Once cointegration is ascertained, we turn to the estimation of the cointegration relationship. There are two ways to estimate it: the Error Correction Model (ECM) approach and the system VAR. We have used ECM as suggested by Engle and Granger (1987), in order to determine whether or not there is a long run relationship among variables. The absence of serial correlation among residuals permits the implementation of the ECM procedure. To overcome simultaneity problems, two stage least square and instrumental variables have been implemented.

A trend, a season and two dummy variables relative to 1995 and 1998 have been included in equation 24. The dummy for 1998 represents the Russian financial crisis which occurred in August 1998. The dummy for 1995 reflects the drop in industrial production which materialized in the last months of the year. Their presence is clear looking at the outliers in the residuals (Figure 1).

Figure 1 Residual Graph



The final ECM specification estimated by LS technique is formalised by:

$$dlogx = \alpha + \beta_1 logx_{t-1} + \beta_2 logp_{t-1} + \beta_3 logy_{t-1} + \beta_4 logw_{t-1} + \chi_2 d\,logp_{t-1} + \chi_3 d\,logy_{t-1} + \gamma_2 d\,logp_{t-2} + $$
$$+ \gamma_4 d\,logw_{t-2} + \lambda_4 d\,logw_{t-4} + trend + D95 + D98 + seas(4) + \upsilon_t \qquad (25)$$

The estimated model is robust in terms of autocorrelation and normality of residuals. In particular, the residuals are white noise, because using the Breusch-Godfrey Serial Correlation LM Test, the null hypothesis (no serial autocorrelation exists) cannot be rejected (Table 5 appendix). Normality of the residuals has been examined performing the Jarque-Bera multivariate test. The fit as measured by R2 is good (0.75). Finally, the t-statistic of lx (-4.66), in accordance with the PPS test, emphasises the existence of a long run cointegration relationship among variables, i.e. variables cannot move independently from each other.

The null hypothesis, Ho: no cointegration vs. H1: cointegration has been rejected since:

tECM = -4.66 < tCRIT –3.98

The long run cointegration relationship among variables is given by the Bewley Transformation (Tab.4 appendix) formalised as follows:

$$log\ x_{t-1} = -\ 2.402 * log\ p_{t-1} - 1.453 * log\ y_{t-1} + 3.315\, log\ w_{t-1} \qquad (26)$$
$$\qquad\qquad (-2.026) \qquad\qquad (-2.805) \qquad\qquad (3.894)$$

The values in brackets are t-values.
The short-run elasticities are reported below:

$$log\ x_{t-1} = -\ 0.23 * log\ p_{t-1} - 1.17 * log\ y_{t-1} + 2.56\, log\ w_{t-1} \qquad (27)$$
$$\qquad\qquad (-2.039) \qquad\qquad (-2.338) \qquad\qquad (2.214)$$

5.3 Testing for Structural Breaks

Over the 90s, Russia experienced the effects of important institutional changes as a consequence of the Soviet Union break down. Moreover, in August 1998 Russia went through a financial crisis, which was mainly triggered by fears on the prospects of political stability and ongoing declines in oil prices. One could then suspect the possibility of structural breaks in our model. However neither the results of the Chow break point test nor the Cumsum test give support to the hypothesis of significant structural breaks.

To carry out the Chow test, our data were partitioned into two sub-samples from May 1994 to July 1998 and from August 1998 to November 2001.

The gauged Chow statistics for August 1998 are reported in Table 2 . Since the estimated F-statistic falls inside the tabulated F-statistic range, we conclude that the null of no structural change cannot be rejected

Table 2 Chow Breakpoint Test: 1998:08

F-statistic	1.086359	Probability	0.386337
Log likelihood ratio	17.89172	Probability	0.161674

Tabulated F-statistics
F(13, 78) = 2.34 at 1% significance
F(13, 78) = 1.92 at 5% significance

Also the log likelihood ratio statistic, based on the comparison of the restricted and unrestricted maximum of the (Gaussian) log likelihood function, does not reject the null of hypothesis of no structural change.

To check whether Russia went through other structural breaks during the considered time frame, the CUSUM test has been performed (Brown, Durbin, and Evans, 1975). The test, based on the cumulative sum of the recursive residuals, does not find parameter instability as the cumulative sum does not go outside the area between the two 5% critical lines.

6 Estimation Results: Discussion

The coefficients of the independent variables (eq. 26) are the price and income elasticity of the Russian exports. Put differently, the coefficients show the sensitivity of Russian exports to changes in relative prices and domestic and world income. In line with the literature, price elasticity and income elasticities enter the final equation with the expected signs and they are highly significant. More specifically, the long run price elasticity is –2.4. A reduction in relative prices of 10%, i.e. a real depreciation of 10%, brings about a rise in exports of 24%. Therefore, the Marshall-Lerner condition, which states that a real depreciation improves the current account if exports are sufficiently elastic to the real exchange rate, is met. The effects of real devaluation are very impressive. Relative prices are hence important determinants of trade flows. A rouble depreciation, in fact, affects significantly trade flows and contributes to ameliorate Russia's current account balance. Russia has experienced a gain in price competitiveness and an increase in export market share since the rouble devaluation in 1998. The sensitivity to prices of Russian exports suggests that careful attention should be paid to price elasticity in order to determine exchange rate movements, which in turn are useful to evaluate the international competitiveness of Russia. It is known that policy makers face an impossible task if they have to conciliate the so called "inconsistent quartet" of (1) free trade, (2) complete international capital mobility, (3) fixed or managed exchange rates and (4) autonomy of monetary policy. Both economic theory (Mundell, 1963; Padoa Schioppa, 2002) and historical experience, in fact, provide a striking confirmation that the four objectives cannot coexist, but one of them must be overlooked. As indicated by the Mundell quartet, if three of the four conditions—free trade, capital mobility and independent monetary policy—are fulfilled, it is not possible to maintain a fixed exchange rate; the latter has to be floating. In this context, while Europe has "reconciliated the inconsistent quartet by moving from autonomous national monetary policies to monetary Union" (Padoa Schioppa, 2002), USA and Japan have opted for free trade, free capital movements and independent monetary policy, China has given away free capital movements, Russia and Thailand still pursue all four objectives (Table 3). The latter strategy is an unsustainable or 'inconsistent' policy mix, which can lead to strong financial crises. Therefore, it is necessary to give away one of the objectives. Since the Russian Federation is looking for free trade and capital mobility and the Central Bank of Russia controls monetary policy, the fixed exchange rate should be given up. The estimation of trade elasticities, as a consequence, becomes extremely important because allows to predict real exchange rate changes necessary to obtain a given change in the current account.

Table 3 The Inconsistent Quartet

Country/Area	Free Trade	Free Capital Movements	Autonomous Monetary Policy	Fixed Exchange Rate
EU Members	yes	yes	no	yes
USA/ Japan	yes	yes	yes	no
China	partially	no	yes	yes
Thailand	yes	yes	yes	yes
Russia	yes	yes	yes	yes

Source: Own Presentation based on IMF World Economic Outlook, various issues

To make plain this last concept consider the exports of Russia, the estimated elasticity and the value of the rouble in 1994. The movement of the Russian exchange rate can be predicted using some simple indices. Russia's exports between 1994 and 2001 have grown from 67.8 to 101.6 billion US dollars (the Central Bank of the Russian Federation, 2003) the real trade weighted exchange rate with base 1995=100 was 105 in 1994 (RET, 2003). Without considering imports and supposing a price elasticity equal to -1, the raise in exports by 49.8% throughout the considered time frame would have required an equal depreciation of the rouble, i.e. the weighted exchange rate would go from 105 to about 52.29. But with an elasticity of -2.4, such export increase needs a real depreciation of about 21%[3], i.e. the weighted exchange rate would have had to change from 105 to about 83[4]. This value actually, matches the value of the real trade exchange rate issued by the Russian economic Trends, 83.7.

The estimated price elasticity of -2.4 refers to exports excluding oil and gas and their products. In general, price elasticity is higher when a country is specialised in low and medium technology goods. Excluding oil and energy products, Russia in fact is specialised in wood, mineral products, fish and fur-skins. By contrast, price elasticity is lower in countries whose trade includes a higher proportion of exported goods offering more opportunities for differentiation, which can be based on factors such as technological sophistication, quality or brand image. In a previous estimation conducted on the Russian export demand inclusive of oil and its products, we found a foreign demand for Russian products quite inelastic to prices (0.79). The reason can be traced back to the fact that the demand for oil and all energy products is not influenced that much by changes in prices. This means that in this latter case, a reduction in prices of 10% would cause a rise in exports of only 8%.

The demand for exports is elastic with respect to worldwide disposable income. The high long run elasticity (3.3) implies that the value of exports increases substantially when world income increases.

3 Note that $\Delta Q/\Delta p$=elasticity, thus $49.8\%/\Delta p=-2.4$ and $\Delta p=20.75\%$

4 105*21%= 22.05 105-22.05=82.95

The demand for exports is significantly linked to domestic disposable income too. The value of exports increases by 14% when domestic income decreases by 10%. A growth in domestic income produces an increase in demand for domestic and foreign products—it pushes imports up and reduces exports. In other words, higher absorption would have been spent on imports. The negative sign obtained from the estimations can be explained by the fact that national GDP has been netted of exports. In another previous estimation we took into account the total GDP with exports. In this last case the sign was, as might be expected, positive.

The higher the elasticity of foreign demand for Russian products to world income, the stronger the exports will be as an engine of growth. The higher the price elasticity, the more competitive is the international market for exports of the particular country, and thus the more effective a real devaluation will be in upgrading export receipts.

The short term elasticities are smaller than the long run ones. In the short period, the effects of competitiveness on exports are quite scanty, while the response of exports to income changes is more immediate. Consequently, in the short term, Russian exports are dominated by movements of worldwide and domestic real income, which become a crucial determinant of economic performance in Russia, while changes in competitiveness take longer to affect export performance.

7 Conclusion

The paper provides price and income elasticities of the export demand function for Russia, estimated within a cointegration ECM framework and taking into account the new empirical approach by Pesaran-Shin-Smith (2001). An extended version of the Imperfect Substitute Model has been implemented. The period of analysis goes from December 1993 to November 2001. Monthly data have been used.

The long-run price and income elasticities have the expected signs and are highly significant. In particular, the long run price elasticity is found to be -2.40. The high price elasticity is explained by the low proportion of high-technology goods exported by Russia. It does emerge that there is a specialisation in products that allow less differentiation and are therefore, subject to higher price-competition from third countries. The long run world income elasticity is equal to 3.31, while the Russian income elasticity is -1.45. In the short period, the effects of competitiveness on exports are limited, while the reaction of exports to income changes is immediate. Thereby, in the short term, Russian exports are dominated by movements of worldwide and domestic real income, which becomes a crucial determinant of economic performance in Russia, while changes in competitiveness take longer to affect export performance.

The cointegration equation is robust in terms of autocorrelation and normality of residuals and it has a high goodness of fit.

References:

Aigner, D.J and S.M. Goldfeld (1974), 'Estimation and Prediction from Aggregate Data when Aggregates are Measured more Accurately than their Components', *Econometrica*, 42, 113-134.

Aquino, A. (1982), 'Crescita Economica e Bilancia dei Pagamenti: un Tentativo di Sintesi', *Giornale degli Economisti e Annali di Economia*, 3, 185-205.

Banco de Espaňa (2003), 'Comparative Analysis of Export Demand for Manufactures in the Euro Area Countries', *Economic Bulletin*, July.

Brown, R.L., J. Durbin, and J.M. Evans (1975), 'Techniques for Testing the Constancy of Regression Relationships Over Time', *Journal of the Royal Statistical Society*, 37, 149–192.

Campbell, J.Y. and P. Perron (1991), 'Pitfalls and Opportunities: What Macroeconomists Should Know About Unit Roots', *NBER Technical Working Paper* , No.100, April.

Chipman, S.J. (1985), 'Estimation of Net-Import Demand Functions for the Federal Republic of Germany, 1959-1982', *Herbert Giersch (editor)*, Probleme und Perspektiven der weltwirtschaftlichen Entwicklung, Jahrestagung des Verein für Socialpolitik, Gesellschaft für Wirtschaftsund Socialwissenschaften in Travemünde vom 17.-19. September 1984, Neue Folge, Band 148, *Berlin: Duncker & Humblot*, 197-213.

Clarida, R. (1994), 'Co-integration, Aggregate Consumption, and Demand for Imports: A Structural Econometric Investigation', *American Economic Review*, 84, 298-308.

Cubadda, G., S. Fachin and F. Nucci (1999), 'Disaggregated Import Demand functions for the Italian Economy', in C. Sardoni and P. Kriesler (eds), *Keynes, Post-Keynesianism and Political Economy*, London, Routledge.

Deardorff A. And R. M. Stern (1978), 'The Terms-of-Trade Effect on Expenditure: Some Evidence from Econometric Models', *Journal of International Economics*, 8, 409-414.

de la Croix, D. and J.P. Urbain (1995), 'Intertemporal Substitution in Import Demand and Habit Formation', *Working Paper at the HCM workshop in Aarhus.*

Engle, R.F. and C.W.J. Granger (1987), 'Co-integration and Error-Correction Representation: Estimation and Testing', *Econometrica*, 55, 251–276.

Faini, R., F. Clavijo and A. Senhadji (1992), 'The Fallacy of Composition Argument: Is it Relevant for LDCs' Manufactures Exports?', *European Economic Review*, 36, 865–882.

Giovannini, A. (1988), 'Exchange Rates and Traded Good Prices', *Journal of International Economics*, 24, 45-68.

Goldstein, M. and M.S Khan (1985), 'Income and Price Effects in Foreign Trade', in Jones RW, Kenen PB (eds.), *Handbook of International Economics* Vol. II, North-Holland.

Hooper, P., K. Johnson and J. Marquez (1998), 'Trade Elasticities for G-7 Countries', *International Finance Discussion Papers*, No.609.

Hung, W., Y. Kim and K. Ohno (1993), 'Pricing Exports: a Cross-Country Study', *Journal of International Money and Finance*, 12, 3-28.

IMF, World Economic Outlook, various issues.

Kalecki, M. (1971), 'On Foreign Trade and Domestic Exports', in *Selected Essays on the Dynamics of the Capitalist Economy 1933-1970*, London, Cambridge University Press.

Kravis, I. B. and Lipsey R.E. (1983), 'Toward an Explanation of National Price Levels', *Princeton Studies in International Finance*, 52, 21-38.

Leamer, Edward E. and Robert M. Stern (1970), *Quantitative International Economics*, Boston, Allyn and Bacon.

Lipsey, R. (1978), 'Price Behaviour in the Light of Balance of Payments Theories', *Journal of International Economics*, 8, 193-246, NBER Reprint No. 26.

Madsen, J. B. (1998), 'Errors-in-Variables, Supply Side Effects, and Price Elasticities in Foreign Trade', *Weltwirtschaftliches Archiv*, 134(4), 612-637.

Marquez, J. and C. McNeilly (1988), 'Income and Price Elasticities for Exports of Developing Countries', *Review of Economics and Statistics*, 70, 306–314.

Mundell, R. (1963), 'Capital mobility and stabilization policy under fixed and flexible exchange rates', *Canadian Journal of Economic and Political Science*, 29, 475-485.

Nielsen, H. B. (2001), 'An I(2) Cointegration Analysis of Price and Quantity Formation in Danish Manufactured Exports', *Discussion Paper No.1 Institute of Economic*s, University of Copenhagen.

Ogaki, M. (1992), 'Engel's Law and Cointegration', *Journal of Political Economy*, 100, 1027-1046.

Orcutt, G.H (1950), 'Measurement of Price Elasticities in International Trade', *The Review of Economics and Statistics*, 32, 117–132.

Ostry, J. and A. Rose (1992), 'An Empirical Evaluation of the Macroeconomic Effects of Tariffs', *Journal of International Money and Finance*, 11, 63–79.

Padoa Schioppa, T. (2002), 'Competition, Co-operation, Public Action: Three Necessary Drivers for European Financial Integration', The European Central Bank, Dinner Speech for the Launching Workshop of the *ECB-CFS Research Network* on "Capital Markets and Financial Integration in Europe" Frankfurt am Main, 29 April.

Pesaran, M. H., Shin, Y. and Smith, R.J. (2001), 'Bounds Testing Approaches to the Analysis of Level Relationships', *Journal of Applied Econometrics*, 16, 289-326.

Prais, S. J. (1962), 'Econometric Research in International Trade: A Review', *Kyklos*, 15, 560-579.

Reinhart, C. (1995), 'Devaluation, Relative Prices, and International Trade', *IMF Staff Papers*, 42, 290-312.

Rose, A. (1990), 'Exchange Rates and Trade Balance: Some Evidence from Developing Countries', *Economic Letters*, 34, 271–275.

Rose, A. (1991), 'The Role of Exchange Rates in Popular Models of International Trade: Does the Marshall-Lerner Condition Hold?', *Journal of International Economics*, 30, 301–316.

Russian Economic Trends (RET), Produced by the *Russian-European Centre for Economic Policy* in cooperation with the Working Center for Economic Reform under the Government of the Russian Federation, various issues.

Senhadji, A. and C. Montenegro (1998), 'Time Series Analysis of Export Demand Equations: A Cross-Country Analysis', *IMF Working Paper*, No.149.

Senhadji, A. and C. Montenegro (1999), 'Time Series Analysis of Export Demand Equations: A Cross-Country Analysis', *IMF Staff Papers*, 46, 3.

Stern R. M., J. Francis and B. Schumacher (1976), *Price Elasticities in International Trade*, London, Macmillan Press.

Taplin, G. B. (1973), 'A Model of World Trade', in R.J. Ball (eds), *The International Linkage of National Economics Models,* Amsterdam: North-Holland.

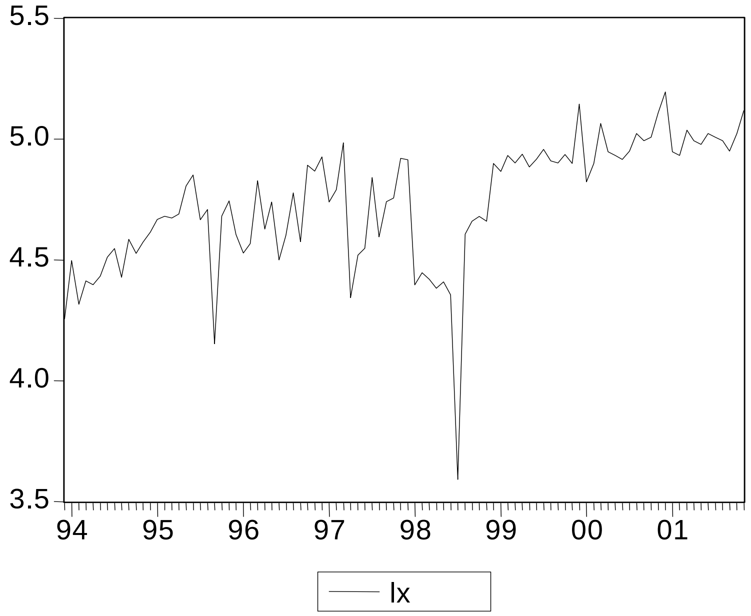
The Central Bank of the Russian Federation, CBR (2003), *Bulletin of Banking Statistics*, Moscow.

Wolf, H., J. De Gregorio and A. Giovannini (1994), 'International Evidence on Tradables and Nontradables Inflation', *European Economic Review*, 32, 1225-1544.

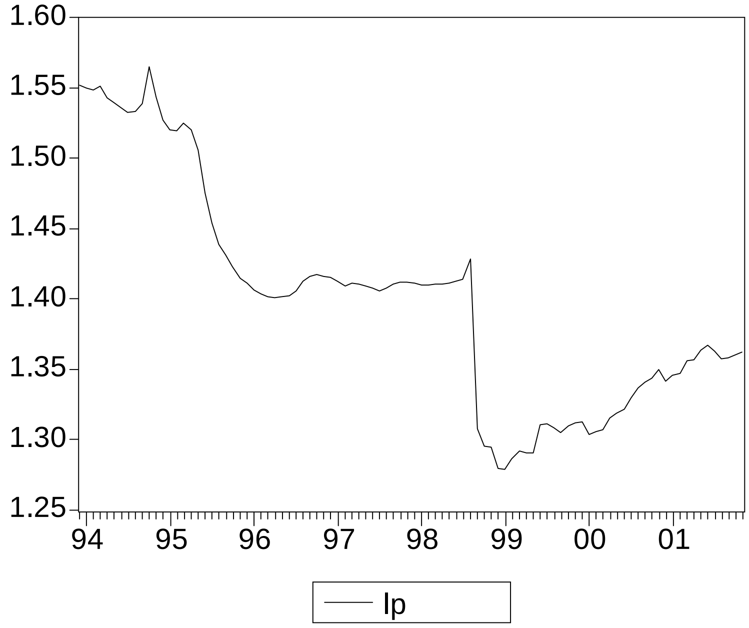
Wolf, H. and J. Haskel (2000), 'From Big-Macs to iMacs: What Do International Price Comparisons Tell Us?', *World Economics*, 1, 167-178.

Appendix

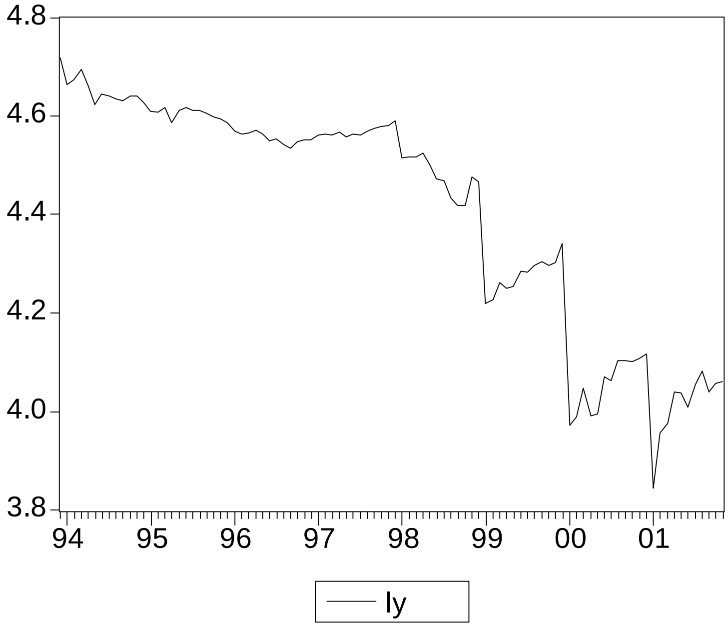
Figure 2 Time Series Dynamics
Russian Exports without oil and petroleum products (log values) 1994-2001



Relative Prices of Russian products in terms of foreign products (log values) 1994-2001



Russian GDP, excluding exported goods (log values) 1994-2001.



OECD industrial production (log values) 1994-2001

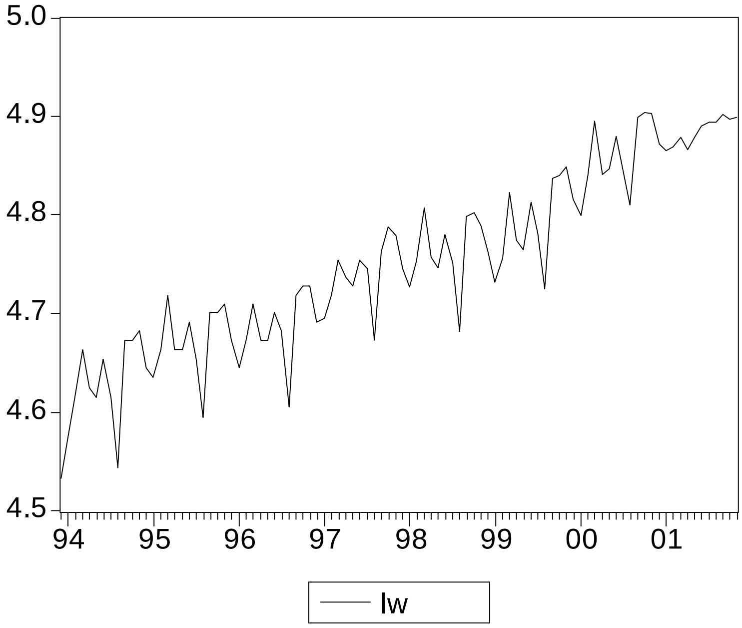


Table 4 Unit Root Tests

	ADF-Test on level	ADF-Test on first difference
logx	-2.803986	-4.866052
logy	-2.244854	-5.320290
logp	-1.672600	-4.069250
logw	-3.336670	-9.687245

On level	1-st differences
1% Critical Value* -4.0613	1% Critical Value* -4.0625
5% Critical Value* -3.4591	5% Critical Value* -3.4597
10% Critical Value* -3.1554	10% Critical Value* -3.1557

	PP-Test on level	PP-Test on first differences
logx	-4.213866	
logy	-0.860524	-11.97013
logp	-1.856028	-8.855690
logw	-2.241063	-16.32586

On level	1-st differences
1% Critical Value* -4.0570	1% Critical Value* -4.0580
5% Critical Value* -3.4571	5% Critical Value* -3.4570
10% Critical Value* -3.1542	10% Critical Value* -3.1545

*MacKinnon critical values for rejection of hypothesis of a unit root.

Table 5 Final ECM Specification

Dependent Variable: dlogx
Method: Least Squares
Sample(adjusted): 1994:05 2001:11
Included observations: 91 after adjusting endpoints

Variable	Coefficient	Std. Error	t-Statistic	Prob.
logx(-1)	-0.390950	0.106682	-4.664631	0.0004
logp(-1)	-0.942510	0.421295	-2.237173	0.0281
logy(-1)	-0.570055	0.172455	-3.305519	0.0014
logw(-1)	1.299598	0.292896	4.437058	0.0000
D95=0	0.515969	0.159690	3.231071	0.0018
@TREND	-0.008856	0.002711	-3.266636	0.0016
dlogw(-2)	-0.720001	0.412753	-1.744385	0.0850
dlogy(-1)	0.458957	0.287627	1.595669	0.1146
@SEAS(4)	-0.238841	0.064049	-3.729055	0.0004
dlogw(-4)	-1.292785	0.408322	-3.166095	0.0022
D98=0	-0.543883	0.200235	-2.716221	0.0081
dlogp(-1)	0.359716	0.225580	1.594628	0.1148
dlogp(-2)	-0.181583	0.215821	-0.841361	0.4027

R-squared	0.759276	Mean dependent var	0.007903
Adjusted R-squared	0.691472	S.D. dependent var	0.213302
S.E. of regression	0.152108	Akaike info criterion	-0.796886
Sum squared resid	1.804678	Schwarz criterion	-0.438192
Log likelihood	49.25831	Durbin-Watson stat	2.157067

Breusch-Godfrey Serial Correlation LM Test:

F-statistic	1.430260	Probability	0.232429
Obs*R-squared	6.530453	Probability	0.162882

Table 6 Bewley Transformation

Dependent Variable: logx
Method: Two-Stage Least Squares
Sample(adjusted): 1994:05 2001:11
Included observations: 91 after adjusting endpoints
Instrument list: logx(-1) logp(-1) logy(-1) logw(-1) D95=0
@TREND dlogw(-2) dlogy(-1) @SEAS(4) dlogw(-4)
D98=0 dlogp(-1) dlogp(-2)

Variable	Coefficient	Std. Error	t-Statistic	Prob.
dlogx	-1.546192	0.693215	-2.230467	0.0286
logp(-1)	-**2.402030**	1.185524	-2.026134	0.0462
logy(-1)	-**1.453184**	0.518119	-2.804733	0.0064
logw(-1)	**3.315035**	0.851247	3.894329	0.0002
D95=0	1.312926	0.573126	2.290815	0.0247
@TREND	-0.022561	0.008779	-2.569727	0.0121
dlogw(-2)	-1.831566	1.205784	-2.118984	0.0328
dlogy(-1)	1.170151	0.760711	2.338234	0.0280
@SEAS(4)	-0.608081	0.234670	-2.591213	0.0114
dlogw(-4)	-3.290312	1.424612	-2.309619	0.0236
D98=0	-1.379639	0.796230	-1.732715	0.0871
dlogp(-1)	0.916194	0.619295	2.039416	0.0431
dlogp(-2)	-0.462801	0.557676	-2.729875	0.0091

R-squared	*0.732160*	Mean dependent var	0.006295
Adjusted R-squared	*0.627063*	S.D. dependent var	0.170567
S.E. of regression	*0.117300*	Akaike info criterion	-1.248990
Sum squared resid	*0.963150*	Schwarz criterion	-0.669561
Log likelihood	*77.82903*	F-statistic	6.015000
Durbin-Watson stat	*2.140877*	Prob(F-statistic)	0.000000

Breusch-Godfrey Serial Correlation LM Test:

Obs*R-squared	6.558747	Probability	0.161126