An empirical analysis of the trade balance in post-communist Albania

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Abstract. In this study we empirically analyze the relationship between real exports and real imports during the post-communist transition in Albania. The country liberalized its trade in the early 1990s after the fall of communism and has seen its trade deficit increase since then. The continuous deterioration of the country’s trade balance can create serious problems for its economy. We employ the ARDL bounds testing approach to a transformed equation derived from the Husted’s (1992) model of inter-temporal budget constraint. The data used are annual observations from 1992 to 2012 obtained from the World Development Indicators 2013. Results show that a long-term relationship exists between real exports and real imports. In addition, we find that imports grow faster than exports and the country’s trade deficit is likely to increase with the increase of imports. More specifically, a 10% increase of real imports is accompanied by a 9.61% increase of real exports. The rising trade deficit should be a cause of concern for the country’s policymakers. They should try to adopt effective export-promoting policies as such policies would be particularly suitable in reducing the trade deficit in Albania.

Keywords: exports, imports, relationship, Albania, ARDL.

JEL Codes: F10, P20.

1. Introduction

Trade balance is an important indicator of an economy’s health because it is closely related to a country’s competitiveness in the world markets. Persistent trade imbalances may indicate low competitiveness and can be a source of economic crisis if not kept under control. Albania has been experiencing a trade deficit since the country opened its economy following the collapse of communism in 1992. During the transition period to the market economy the trade deficit has kept growing. The Bank of Albania data show that it has been growing at an average rate of 10.33% per year during the period 1993-2012. There has been a growing debate among researchers on the ways to address the country’s growing trade imbalance. The general belief is that the continuous widening of the trade balance can become a serious problem for a small open economy like Albania and important structural reforms need be undertaken now. The underlying theoretical argument is that the country must rely on a larger amount of foreign capital to finance the resulting extra investment. The increasing reliance on foreign sources of finance can manifest itself in higher interest rates paid by the country to borrow money in the world markets, as creditors require higher interest rates to compensate for the higher risk of default. This translates into higher interest payments.
and a larger stock of debt which calls for further increase in the borrowing interest rate. The heavy debt burden can seriously compromise the economic welfare of the coming generations both through higher debt payments and through its potential adverse impact on the country’s economy ability to grow. As Arize (2002) emphasizes, understanding the long-term relation between exports and imports in a particular country is crucial for formulating appropriate macroeconomic policies to balance trade in that country.

However, despite the large number of empirical research on this topic the results are far from being unanimous. Researchers report contradictory findings even for the same country. For instance, Arize (2002), Upender (2007), Khokhar (2010) and Ramakrishna (2013) all find evidence of cointegration between exports and imports in India while Konya and Singh (2008) do not find evidence of a long-term relation between these two variables.

Empirical studies on this topic for Albania are very limited. As Albania tries to become a member of EU its ability to compete in the EU market has become a central issue of concern among scholars and policymakers. Studying the behaviour of its trade balance can provide useful insights regarding the country’s competitiveness thus helping to formulate appropriate economic policies that can bring the country nearer to the EU. We try to provide additional evidence on the relationship between exports and imports in Albania using more recent data that cover a longer period of time than those used by Çil Yavuz and Kiran (2012). To our knowledge it is the only study that analyzes this relationship in Albania. In contrast to Çil Yavuz and Kiran (2012) that use quarterly data we use annual data. Tang and Alias (2005) believe that the use of annual data is more appropriate in assessing cointegration than the use of quarterly data as the seasonality of the latter can negatively impact the reliability of the estimated long-run coefficients. Due to short time series we assess this relation employing an alternative econometric method that is particularly appropriate when dealing with small samples like ours.

The rest of this study is organized as follows. Section 2 summarizes the existing empirical literature on this topic. The following section presents briefly the econometric model and the data used in this study. Section 4 summarizes the econometric methodology employed. Section 5 presents the estimated results and the final section concludes.

2. Existing empirical evidence on the export-import relation

Choong et al. (2004) is one of the earliest empirical studies on the relationship between exports and imports that belong to the last decade. The authors investigate this relation for Malaysia using annual IMF data during the period 1959-2000. Like almost all the other empirical studies that follow the study employs an econometric model that is based on Husted (1992). Husted’s model is central to the study of current account balance. It allows assessing the effectiveness of the adopted macroeconomic policies in a country by checking if the country satisfies its inter-temporal budget constraint. To this end, Choong et al. (2004) use Johansen and Juselius multivariate cointegration test on 4 different measures for each of the two variables of interest. The four measures consist of both nominal and real values in terms of the domestic currency and the USD. Results indicate that exports and imports converge in the long run which the authors interpret as evidence in favour of the effectiveness of the government policies. Tang (2005) tries to provide additional evidence for this relation in Malaysia. The annual data used cover the period 1960-2001 and are collected from the World Bank. They are deflated using the export price deflator and import price deflator to derive the respective real values and then transformed in natural logarithm forms. Johansen’s cointegration test results suggest no cointegration between exports and imports contradicting the previous study’s finding. Rahman (2011) in a more recent study using two alternative cointegration tests, namely Engle-Granger and Johansen tests, finds evidence in favour of export-import cointegration in Malaysia but fails to find such a
relation in Indonesia. Tang and Alias (2005) also find mixed results when studying this relation individually for 27 Organization of Islamic Conferences member countries. They employ the Engle and Granger cointegration method on real annual data transformed into natural logarithms. Herzer and Nowak-Lehman (2005) employ the Gregory and Hansen (1996) cointegration method to study the relationship between Chilean exports and imports. The authors use annual data in constant local currency transformed in natural logarithms. The data cover the period 1975-2004 and are obtained from the Central Bank of Chile. Results suggest a cointegrating relationship between the two variables of interest. Perera and Varma (2008) apply the same cointegration method on annual export and import data for Sri Lanka from 1950 to 2006. The time series are expressed both in current domestic currency and in USD. Most of the evidence indicates lack of cointegration between the two variables. Jain and Sami (2012) explore this relation for Singapore using time series data from 1976 to 2009 obtained from the World Development Indicators. The data are in local currency and are transformed into natural logarithms. Three alternative cointegration methods are employed on the econometric model that follows Husted (1992), in order to increase the reliability of results. They consist of Autoregressive Distributed Lag (ARDL) bounds testing approach of Pesaran et al. (2001), Dynamic Ordinary Least Squares (DOLS) of Stock and Watson (1993) and Fully Modified Ordinary Least Squares (FMOLS) of Phillip and Hansen (1990). Results support cointegration between exports and imports in Singapore. There are a number of studies on the relation between exports and imports in India. Upender (2007) uses Engle and Granger cointegration methodology on annual time series data for exports and imports in nominal terms. The data are collected from official national sources and cover the period 1949/50-2004/05. They are transformed in natural logarithms and checked for stationarity before estimation. Results indicate that exports and imports in India are cointegrated. Konya and Singh (2008) repeat the exercise using annual data in nominal terms for the same period expressed in both domestic currency and USD. Using two alternative cointegration tests on the natural logarithms of the variables the authors do not find evidence of cointegration between India’s exports and imports. Khokhar (2010) joins this debate for India. The author uses two alternative cointegration tests that include ARDL and Engle-Granger on annual data on exports and imports in real domestic currency from 1960/61 to 2006/07. The author reaches similar conclusions to Upender (2007). Ramakrishna (2013) provides further evidence in favour of a long-run equilibrium relation between exports and imports in India using both the Engle-Granger and Johansen cointegration methods. The annual data used in his analysis are obtained from the World Development Indicators for the period 1970-2010. They are in nominal USD and are transformed in natural logarithms. The findings for Iran are similar. Çelik (2011) studies this relationship for Turkey employing Engle-Granger cointegration method on monthly export and import data in real terms from 1990 to 2010. Results suggest a long-run relationship between these two variables that widens the trade imbalance. Şonjeet al. (2010) investigates this relation individually for 16 former communist European economies applying the Johansen approach on quarterly data. The data are obtained from the IMF and are transformed in natural logarithms. The starting years of the time series variables are not the same for all countries due to data availability, while their last observations corresponds to the end of year 2006. No general pattern emerges as results are mixed depending on the particular country considered. Çil Yavuz and Kiran (2012) analyze sustainability of current account deficit in 5 ex-communist European countries (including Albania) employing a more recent cointegrating test that allows for the presence of two endogenously determined structural breaks. The cointegration test used has been developed by Hatemi-J (2008) and is an extension of Gregory and Hansen (1996) approach that considers one potential unknown structural break. The data on exports and imports for each transition country considered are quarterly and cover the period 1995:01 to 2011:03. The data source is IMF. As the focus is on current account sustainability not just trade balance sustainability, imports represent a wide measure that consists of imports of goods and services plus net interest payments plus net transfer payments. The series are converted
in real terms prior to checking for their stationarity. Results suggest cointegration in each of the 5 transition countries.

3. Model and data

Based on Husted (1992), we employ the following basic econometric model:

\[ \ln\text{expo}_t = \alpha + \beta \ln\text{impo}_t + \varepsilon_t \]  

where \( \ln\text{expo}_t \) is the natural logarithm of the exports of goods and services in constant 2005 USD, \( \ln\text{impo}_t \) is the natural logarithm of the imports of goods and services in constant 2005 USD, \( t \) denotes time, \( \alpha \) is a constant term and \( \varepsilon \) is the white noise error term. The data are annual and are obtained from the World Development Indicators (2013). The variables are transformed by the “natural” logarithm because this transformation facilitates result interpretation. More specifically, the first differences of the transformed variables can be interpreted as the growth rate of the respective variables before transformation.

Prior to estimation we check the nature of the 2 time series variables under consideration.

4. Estimation method

We use 4 traditional unit root tests that include Augmented Dickey-Fuller (ADF), Phillips-Perron (PP), Kwiatkowski-Phillips-Schmidt-Shin (KPSS) and Dickey-Fuller Generalized Least Squares (DF-GLS.) The test equations are chosen based on the visual inspection of the graphs of the variables in levels and their first differences. Both a constant and a trend are included in the test equations when checking the stationarity (nonstationarity) of the variables in levels. In contrast, the test equations for their first differences include a constant only. The unit root test results are presented below.

Table 1. Unit root test results of the time series variables

<table>
<thead>
<tr>
<th>Variable</th>
<th>Level</th>
<th>ADF</th>
<th>PP</th>
<th>KPSS</th>
<th>Result</th>
<th>First difference</th>
<th>ADF</th>
<th>PP</th>
<th>KPSS</th>
<th>Result</th>
</tr>
</thead>
<tbody>
<tr>
<td>lnEXPO_t</td>
<td>-0.254031</td>
<td>-2.055401</td>
<td>0.144864</td>
<td>inconclusive</td>
<td>-3.653202**</td>
<td>-4.83933*</td>
<td>0.292409</td>
<td>stationary</td>
<td></td>
<td></td>
</tr>
<tr>
<td>lnIMPO_t</td>
<td>-1.075234</td>
<td>-2.055401</td>
<td>0.124073</td>
<td>inconclusive</td>
<td>-4.097257*</td>
<td>-4.097257*</td>
<td>0.198927</td>
<td>stationary</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

* and ** indicate rejection of the null hypothesis (\( H_0 \)) for the ADF, PP and KPSS tests at 1% and 5% level of significance respectively.

Notes: Test statistics and critical values are computed by the statistical software EViews 7.1. Lag lengths or bandwidths were automatically selected by the program. Selection of lag lengths for both ADF and PP tests was based on Schwarz Information Criterion (SIC). In the KPSS test Newey-West Bandwidth was selected using Bartlett kernel spectral estimation method. The Critical values in both the ADF and PP tests refer to critical values computed by Mac Kinnon (1996) while those in the KPSS test refer to Kwiatkowski-Phillips-Schmidt-Shin (1992, Table 1). Inference is made based on the most frequently used 1% and 5% levels of significance.

Further evidence regarding the nature of time series data is provided by the DF-GLS unit root test, which is more powerful than the ADF unit root test (Elliot et al. 1996). Like in the ADF test the null hypothesis in the DF-GLS test is that the variable under consideration is non-stationary against the alternative that it is stationary. Its results are presented in the table below.
Table 2. Results of the DF-GLS unit root test for the time series variables

<table>
<thead>
<tr>
<th>Variable</th>
<th>Level</th>
<th>First difference</th>
</tr>
</thead>
<tbody>
<tr>
<td>lnEXPO_t</td>
<td>-2.106368</td>
<td>nonstationary</td>
</tr>
<tr>
<td>lnIMPO_t</td>
<td>-1.439935</td>
<td>nonstationary</td>
</tr>
</tbody>
</table>

* and ** indicate rejection of the null hypothesis (H0) for the DF-GLS test at 1% and 5% level of significance respectively.

Notes: Lag lengths are selected automatically by the software EViews 7.1 using SIC. The critical values in this unit root test refer to the Mac Kinnon (1996) critical values.

Most of the evidence indicates that both lnexpo_t and lnimpo_t are I(1).

Next, we check for cointegration between these two variables. The choice of the cointegration method is based on the unit root test results. The order of integration of the variables (none with order of integration higher than 2), the short span of the time series variables under consideration and its ability to deal with potential omitted variable bias (Narayan 2004) strongly suggest that ARDL is the most appropriate cointegration technique to be employed.

Country-specific empirical studies that employ the ARDL method on small samples of time series data consisting of 20 or less observations are not uncommon (e.g. Pattichis 1999; Gaikwad and Fathipour 2013). ARDL bounds test approach to cointegration is a very popular econometric technique developed by Pesaran et al. 2001. The ARDL model is a dynamic specification which includes lagged values of the dependent and explanatory variables as well as contemporaneous values of explanatory variables to estimate both long and short run relations among several variables of interest.

According to Choong (2005), who summarize Pesaran et al. (2001), the ARDL model can be expressed as a VAR model of order p:

\[ z_t = c_0 + \alpha t + \sum_{i=1}^{p} \eta_i z_{t-i} + \epsilon_t \]  

(2)

where 

\( z_t \) is a column vector of variables \( y_t \) and \( x_t \). \( y_t \) is the dependent variable while \( x_t \) is a column vector of k explanatory variables.

\( c_0 \) represents a \((k + 1)\)-component column vector of intercepts,

\( \alpha \) represents a \((k + 1)\) component column vector of trend coefficients,

\( \eta_i \) represents a \((k + 1) \times (k + 1)\) matrix of VAR parameters for lag \( i \),

\( \epsilon_t \) is a \((k + 1)\)-component column vector of white noise error terms,

\( t \) represents time while \( p \) is the optimal lag length.

It can be written as a VEC (Vector Error Correction) model in the following form:

\[ \Delta z_t = c_0 + \alpha t + \lambda \Delta z_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta y_{t-i} + \sum_{i=0}^{p-1} \Phi_i \Delta x_{t-i} + \epsilon_t \]  

(3)

where 

\( \Delta \) is the first difference operator,

\( \lambda \) is a \((k + 1) \times (k + 1)\) long-run multiplier matrix, and

\( \Gamma_i \) and \( \Phi_i \) are a \((k + 1)\) column vector and a \((k + 1) \times k\) matrix of short-run coefficients respectively.

The long-run multiplier matrix can be partitioned as
Our basic econometric model can be expressed in the ARDL form following the assumptions made by Pesaran et al. (2001) for case v (unrestricted intercepts and unrestricted trends). The restriction $\lambda_{xy} = 0$ should be imposed so that at most a unique long-run relationship between $y_t$ and the regressors be examined. It is shown as the following UECM (unrestricted error correction model):

$$\Delta \ln \text{expo}_t = \beta_0 + \beta_1 t + \beta_2 \ln \text{expo}_{t-1} + \beta_3 \ln \text{impo}_{t-1} + \beta_4 \text{DUM} + \sum_{i=0}^{q} a_i \Delta \ln \text{expo}_{t-i} + \sum_{i=0}^{p} b_i \Delta \ln \text{impo}_{t-i} + \varepsilon_t$$  

(4)

DUM is a crisis dummy included in the model to account for the negative impact of the 1992, 1997 crises on the country’s economy. It takes the value 1 for these two years and 0 otherwise. This equation can also be viewed as an ARDL model of order $(p, q)$. The inclusion of the “one-zero” dummy variable in the model above doesn’t cause any trouble as the fraction of observations where the dummy is nonzero is $2/21 = 9.5\%$ (Pesaran et al. (2001). Choong et al. (2005) also use one such dummy variable in the ARDL model employed to account for the impact of the East Asian financial crisis on growth in Malaysia.

5. Results

Several versions of the dynamic specification are estimated in EViews 7.1 statistical package. The estimated coefficients of the most appropriate specification are presented below.

<table>
<thead>
<tr>
<th>Variable</th>
<th>Coefficient</th>
<th>Std. Error</th>
<th>t-Statistic</th>
<th>Prob.</th>
</tr>
</thead>
<tbody>
<tr>
<td>C</td>
<td>0.080997</td>
<td>1.283463</td>
<td>0.063108</td>
<td>0.9505</td>
</tr>
<tr>
<td>LNEPO(-1)</td>
<td>-0.568615</td>
<td>0.189785</td>
<td>-2.996094</td>
<td>0.0090</td>
</tr>
<tr>
<td>LNIIMPO(-1)</td>
<td>0.546161</td>
<td>0.222915</td>
<td>2.450083</td>
<td>0.0270</td>
</tr>
<tr>
<td>DUM</td>
<td>-0.458812</td>
<td>0.175596</td>
<td>-2.612879</td>
<td>0.0196</td>
</tr>
<tr>
<td>DLNIIMPO</td>
<td>0.632158</td>
<td>0.293732</td>
<td>2.152158</td>
<td>0.0481</td>
</tr>
<tr>
<td>R-squared</td>
<td>0.691764</td>
<td>Mean dependent var</td>
<td>0.131269</td>
<td></td>
</tr>
<tr>
<td>Adjusted R-squared</td>
<td>0.609567</td>
<td>S.D. dependent var</td>
<td>0.223677</td>
<td></td>
</tr>
<tr>
<td>S.E. of regression</td>
<td>0.139764</td>
<td>Akaike info criterion</td>
<td>-0.885411</td>
<td></td>
</tr>
<tr>
<td>Sum squared resid</td>
<td>0.293008</td>
<td>Schwarz criterion</td>
<td>-0.636478</td>
<td></td>
</tr>
<tr>
<td>Log likelihood</td>
<td>13.85411</td>
<td>Hannan-Quinn criter.</td>
<td>-0.836817</td>
<td></td>
</tr>
<tr>
<td>F-statistic</td>
<td>8.415986</td>
<td>Durbin-Watson stat</td>
<td>1.971102</td>
<td></td>
</tr>
<tr>
<td>Prob(F-statistic)</td>
<td>0.000908</td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Fig. 1: Estimated coefficients of the dynamic specification

A summary of the results of the model diagnostic tests performed in EViews 7.1 is presented below.
Table 3. Model Diagnostic Checking

<table>
<thead>
<tr>
<th></th>
<th>Value</th>
<th>p-value</th>
</tr>
</thead>
<tbody>
<tr>
<td>AR(1)</td>
<td>0.005101</td>
<td>0.9431</td>
</tr>
<tr>
<td>JB</td>
<td>3.809962</td>
<td>0.148825</td>
</tr>
<tr>
<td>ARCH(1)</td>
<td>0.022628</td>
<td>0.8804</td>
</tr>
<tr>
<td>RESET(2)</td>
<td>1.337530</td>
<td>0.2963</td>
</tr>
</tbody>
</table>

AR, JB, ARCH and RESET stand for the Breusch-Godfrey serial correlation test, the Jarque-Bera normality test, the ARCH test and the Ramsey’s RESET test respectively. The numbers in brackets represent the number of lags = 1 and number of fitted terms = 2 included in the Breusch-Godfrey serial correlation test, ARCH test and RESET test respectively. The probabilities of the calculated test statistics are shown in square brackets. The results above indicate that the estimated model does not seem to have any serious diagnostic problems such as serial correlation, ARCH effects, non-normality of the residuals and misspecification. (Similar conclusions are derived even when the number of fitted terms included in the RESET is 1).

In addition, the plots of both CUSUM and CUSUM of Squares Tests that are used to check parameter stability suggest that the model is stable during the sample period. They are provided below.

Fig. 2: CUSUM Test.

Fig. 3: CUSUM of Squares Test.
Thus, the estimated coefficients of this specification are valid for interpretation. The explanatory power of this model is high as indicated by the coefficient of determination ($R^2 = 0.691764$). The value of this coefficient implies that the independent variables included in the regression model explain more than 69% of the variation in the dependent variable. Furthermore, the coefficients of all the explanatory variables in levels and their first differences are statistically significant at the most common levels of significance. The “crisis” dummy is also highly significant carrying the expected negative sign. In contrast, the intercept is not relevant.

The presence of cointegration between $\ln \text{expo}_t$ and $\ln \text{impo}_t$ is checked by using the Wald test. The computed $F$-statistic is compared to the critical values of the $F$-statistic provided by Narayan (2005), to decide whether a long-run relationship between these two variables exists.

### Table 4. Critical values for the cointegration analysis provided by Narayan (2005)

<table>
<thead>
<tr>
<th>Significance level ($\alpha$)</th>
<th>Lower bound critical value</th>
<th>Upper bound critical value</th>
</tr>
</thead>
<tbody>
<tr>
<td>1%</td>
<td>8.170</td>
<td>9.285</td>
</tr>
<tr>
<td>5%</td>
<td>5.395</td>
<td>6.350</td>
</tr>
<tr>
<td>10%</td>
<td>4.290</td>
<td>5.080</td>
</tr>
</tbody>
</table>

Note: critical values are cited from Narayan (2005) for case III (unrestricted intercept and no trend) for number of regressors ($k$) = 1 and number of time periods ($n$) = 30. The number of time periods in our study is 21 but we use the critical values computed for a sample size of 30 because it is the smallest sample size for which they are calculated by the author. However, this should have only a small impact on the results.

The computed $F$-statistic is 7.157660. As it is larger than the upper bound critical value of 6.350 at the 5% level of significance the null of no cointegration is rejected implying a long-run relationship between $\ln \text{expo}_t$ and $\ln \text{impo}_t$.

This long-run relation is written as:

$$ln \text{expo}_t = 0.080997 + 0.961ln \text{impo}_t$$  \hspace{1cm} (5)

The long-run coefficients are calculated as suggested by Bardsen (1989). The long-run coefficient of $\ln \text{impo}_t$ is statistically significant at the 5% significance level. Evidence suggests a strong positive relation between the two variables of interest. A 10% increase of imports of goods and services seems to be accompanied with a 9.61% increase of exports of goods and services in Albania. In addition, the Wald test shows that the long-run coefficient is significantly smaller than one.

Results suggest a statistically significant positive relation between exports and imports even in the short run. The estimated short-run elasticity of 0.63 is also statistically significant at the 5% significance level.

### 6. Conclusions

We find evidence in favour of a strong positive long-term relation between real exports and real imports in the post-communist Albania. However, results suggest that imports growth faster than exports and trade imbalance in Albania will continue to grow if this trend is not reversed. The continuous widening of the trade balance can be a source of instability for the country’s economy. What is most important, it can also signal low competitiveness of the country. This means that low competitiveness of Albania’s products in the world markets may be the root cause of its trade imbalance and important structural reforms are necessary. The most appropriate course of action would be to adopt policies that promote the country’s exports by increasing its competitiveness.
7. References


