Likelihood-Based Panel Cointegration Approach to Exchange Rate Sensitivity of Bilateral Agricultural Trade Flows: Turkey Versus Her Major Trading Partners

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Abstract

The purpose of this study is to analyze the exchange rate sensitivity of Turkish agricultural trade flows on bilateral basis between Turkey and her six major trading partners in the European Union for the period 1969-2005. To accomplish this purpose, we apply the likelihood-based panel cointegration analysis recently developed by Larsson and Lyhagen (1999) and Larsson et al. (2001). Results show that the exchange rate is not a major factor effecting Turkish agricultural trade flows in the long-run. We also find that the income levels have significant impacts on Turkish agricultural trade flows.

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1. Introduction

The relationship between exchange rate and agricultural sector as well as agricultural trade flows has been a major area of interest (for example, Schuh, 1974; Chambers, 1981; Batten and Belongia, 1986; Bessler and Babula, 1987; Bradshaw and Orden, 1990; Tegene, 1990; Baek and Koo, 2006; Fidan, 2006). The main drawback associated with these studies is the employment of aggregate trade data, causing the “aggregation bias” problem that an insignificant trade elasticity with one trading partner could offset a significant trade elasticity with another partner (Bahmani-Oskooee and Goswami, 2004). To avoid the aggregation bias problem, recent studies have focused on estimating trade elasticities on bilateral basis (for example, Bahmani-Oskooee and Goswami, 2004; Bahmani-Oskooee et al., 2005; Bahmani-Oskooee and Harvey, 2006; Irandoust et al., 2006). In the agricultural trade literature, on the other hand, researchers have been paid relatively little attention to examine the impacts of exchange rate changes on bilateral agricultural trade flows (for example, Kim et al., 2004; Baek, 2007).

The European Union (EU) has the biggest share of Turkish agricultural trade flows with the world, accounting for more than 40 percent in the exports and close to 20 percent in the imports (TurkStat, 2007). The main destinations of Turkish agricultural exports to the EU are Belgium, France, Germany, Italy, Netherlands, United Kingdom, and Spain which have more than 85 percent of Turkish agricultural exports to and more than 75 percent of Turkish agricultural imports from the EU during the last decade (UMFT, 2006). Fluctuations in Turkish agricultural exports to and imports from these countries will therefore have important impacts on Turkish agricultural trade performance due to the fact of their importance in Turkish agricultural trade flows. However, identifying the exchange rate sensitivity of Turkish agricultural trade flows with these countries has been neglected so far. Although Fidan (2006) recently analyzed the impact of real effective exchange rate on Turkish agricultural exports and imports employing aggregate trade data, to our best knowledge the elasticities of Turkish agricultural trade flows on bilateral basis have not been estimated yet.

The purpose of this study is therefore to analyze the exchange rate sensitivity of Turkish agricultural trade flows on bilateral basis between Turkey and her six major trading partners (Belgium, France, Germany, Italy, Netherlands, and United Kingdom).
in the EU for the period 1969-2005. To this end, we use the likelihood-based panel cointegration analysis recently developed by Larsson and Lyhagen (1999) and Larsson et al. (2001). Irlandoust et al. (2006) estimate bilateral trade equations for Swedish trade data using the likelihood-panel cointegration approach. To our literature knowledge the likelihood-based panel cointegration approach to estimate the elasticities of Turkish agricultural trade flows has not been applied yet.

The paper is organized as follows. The model and econometric methodology are outlined in the next section. The data and the variables are discussed in Section 3. Empirical results and conclusion are represented in Section 4 and 5, respectively.

2. The Models and the Method
The traditional way of analyzing the impact of exchange rate on trade flows is to estimate bilateral export and import volume equations. This approach requires using import and export price indexes in order to obtain real exports and imports. However, this method introduces a bias in estimating price elasticities due to the fact that a country trades different goods with different trading partners (Bahmani-Oskooee and Harvey, 2006). Bahmani-Oskooee and Goswami (2004) developed a direct method of assessing the impact of currency depreciation on bilateral trade flows. They propose export and import equations including export and import values as dependent variable because import and export prices are not available on bilateral basis. Following Bahmani-Oskooee and Goswami (2004), we formulate long-run bilateral agricultural export ad import equations as follows:

\[ \ln \text{EXP}_t = \alpha_0 + \alpha_1 \ln Y_{FR,t} + \alpha_2 \ln \text{EXR}_t + \varepsilon_t \]  
\[ \ln \text{IMP}_t = \beta_0 + \beta_1 \ln Y_{TR,t} + \beta_2 \ln \text{EXR}_t + \omega_t \]  

where \( \text{EXP}_t \) is the value of Turkish agricultural exports to trading partner \( i \), \( \text{IMP}_t \) is the value of Turkish agricultural imports from trading partner \( i \), \( Y_{FR} \) is the real income of trading partner \( i \), \( Y_{TR} \) is the real income of Turkey, and \( \text{EXR}_t \) is the real bilateral exchange rate between Turkey and trading partner \( i \).

In Equation (1), an estimate of \( \alpha_1 \) is expected to be positive due to the fact that economic growth in trading partner’s economy should improve the demand for Turkish agricultural products. Following the same idea, an estimate of \( \beta_1 \) in Equation (2) is
expected to be positive, since an increase in Turkish economy should increase the
demand for agricultural products from a trading partner. However, Magee (1973) has
argued that as the real domestic income increases, the production of importable goods
also increases faster than consumption, reducing the volume of imports and thereby
yielding a negative estimate for the real domestic income. If the real depreciation of
Turkish lira increases agricultural exports and decreases agricultural imports, thus
improving trade balance, $\alpha_2$ is expected to be positive and $\beta_2$ is expected to be
negative.

The cointegration technique is an appropriate methodology to analyze the long-
run relationships among the variables. If long time-series data are not available, such as
our case, one solution of this problem is to apply a panel cointegration test. Various
panel cointegration tests are developed by McCoskey and Kao (1998), Kao (1999), and
Pedroni (1999). However, the main drawback of these tests is to assume a priori a
unique cointegration vector (Asteriou and Hall, 2007: 375). To overcome this problem,
Larsson et al. (2001) developed a panel test for the existence of a common cointegration
rank in panel data based on the likelihood inference for vector autoregressive model
developed by Johansen (1988, 1991, 1995). This approach, contrary to above panel
cointegration tests, allows us to test for cointegration vectors in heterogeneous panel
data models (Asteriou and Hall, 2007: 375).

The cointegration analysis first requires determining the order of integration of
the variables. For this pre-requisite, we use two panel unit root tests developed by
Levin et al. (2002, LL) and Im et al. (2003, IPS). If the variables are non-stationary in
levels and integrated at most $I(1)$, we then carry out the panel cointegration analysis.

We use the likelihood-based panel cointegration analysis recently developed by
Larsson and Lyhagen (1999) and Larsson et al. (2001). Consider a panel data set that
consists of $N$ cross-section observed over $T$ time periods. Let $i = 1, ..., N$ index the
groups, $t = 1, ..., T$ the sample period and $j = 1, ..., p$ the variables in each group. The
observed p-vector for group $i$ at time $t$ is given by $Y_{it} = (y_{i1t}, y_{i2t}, ..., y_{ipt})'$ where $y_{ijt}$
denotes the $i$th group, the $j$th variable at time $t$. Note that the number of time-series
observations can vary between the groups, but for notational convenience a common $T$
is used for the presentation. For simplicity, no deterministic components are assumed.
Suppose that the data generating process for each cross-section can be represented by the following heterogeneous VAR \((k_i)\) model:

\[
Y_{it} = \sum_{k=1}^{k_i} \Pi_{ik} Y_{i,t-k} + u_{it} \quad i = 1, \ldots, N
\]

(3)

where for each group \(i\) the values \(Y_{i,t-k+1}, \ldots, Y_{i,0}\) are considered fixed and the errors \(u_{it}\) are independent identically distributed \(u_{it} \sim N_p(0, \Omega)\). Then heterogeneous error correction model can be written as:

\[
\Delta Y_{it} = \Pi_{i} Y_{i,t-1} + \sum_{k=1}^{k_i} \Gamma_{ik} \Delta Y_{i,t-k} + u_{it} \quad i = 1, \ldots, N
\]

(4)

where \(\Pi_i\) is the order of \(p \times p\). If \(\Pi_i\) is of reduced rank, it is possible to write \(\Pi_i = \alpha_i \beta_i\) where \(\alpha_i\) and \(\beta_i\) are of \(p \times r_i\) and full column rank. Note that \(T\) must be large enough so that model (4) can be estimated separately for each cross-section.

The hypothesis considered is that all of the matrices \(\Pi_i, i = 1, \ldots, N\), have rank \(\leq r\). In other words, the hypothesis is that all of \(N\) groups in the panel have at most \(r\) cointegration relationship among the \(p\) variables. The rank hypothesis is stated as:

\(H_0 : \text{rank}(\Pi_i) = r_i \leq r, \text{ for all } i = 1, \ldots, N,\)

\(H_1 : \text{rank}(\Pi_i) = p, \text{ for all } i = 1, \ldots, N,\)

The standard trace test statistic of cointegration rank for group \(i\) is given by

\[
LR_{it} \left( H(r) \big| H(p) \right) = -2 \ln Q_{it} \left( H(r) \big| H(p) \right)
\]

(5)

and the LR-bar statistic is defined as the average of \(N\) individual trace statistics:

\[
\overline{LR}_{NT} \left( H(r) \big| H(p) \right) = \frac{1}{N} \sum_{i=1}^{N} LR_{it} \left( H(r) \big| H(p) \right)
\]

(6)

The statistic proposed as the panel cointegration rank test is the standardized LR-bar statistic, defined by

\[
\gamma_{TR} \left( H(r) \big| H(p) \right) = \frac{\sqrt{N} \left( \overline{LR}_{NT} \left( H(r) \big| H(p) \right) - E(Z_k) \right)}{\sqrt{Var(Z_{ki})}}
\]

(7)

where \(E(Z_k)\) and \(Var(Z_k)\) are the mean and the variance of the asymptotic trace statistic. Asymptotic values of \(E(Z_k)\) and \(Var(Z_k)\) are provided in Larsson et al. (2001) for the case of no deterministic terms in the underlying VAR model. However, Breitung
(2005) has tabulated asymptotic values of $E(Z_k)$ and $Var(Z_k)$ considering various deterministic components in the underlying VAR models. It is important to note that the panel data with large time series dimension relative to cross-sectional dimension are required to avoid serious size-distortions of the panel test.

The proposed testing procedure is the sequential procedure suggested by Johansen (1988). First, $r = 0$ is tested. If the hypothesis is rejected, then $r = 1$ is tested. This sequential procedure is continued until the null is not rejected or the hypothesis $r = p - 1$ is rejected. This procedure gives the rank estimate $r$.

One of the important issues to consider before performing the panel rank test is to decide optimal lag order of the VAR($k$) for each cross-section. The most common procedure in choosing optimal lag length is to estimate a VAR model in levels for a large number of lags and then reduce number of lags one by one to check for optimal value of an information criterion (Asteriou and Hall, 2007). Lutkepohl (1985) shows that the criterion preferred in small samples is the Schwarz Bayesian Criterion (SBC).

Another important issue in the formulation of the dynamic model is to decide whether the VAR model should include an intercept and/or a deterministic trend. Although five distinct models are available, a common practice is to investigate only three of them because the most restrictive specifications are unrealistic (Irandoust et al., 2006). Following Irandoust et al. (2006), we consider the following versions of the deterministic components: restricted intercept and no trend (Model 1), unrestricted intercept and no trend (Model 2), and unrestricted intercept and restricted trend (Model 3). In order to choose the correct model specification, Johansen (1992) proposes to apply the so-called Pantula principle. This principle first involves estimation of all three models and presentation of the standardized LR-bar statistics from the most restrictive model (Model 1) through the least restrictive model (Model 3). Second step comprises moving from Model 1 to Model 3, comparing the test statistic to its critical value at each stage, stopping when null hypothesis of no cointegration is not rejected for the first time (Asteriou and Hall, 2007: 324).
3. Data

We employ annual data covering the period 1969-2005. The trading partners considered are Belgium, France, Germany, Italy, Netherlands, and United Kingdom\(^1\). The description of the variables and their data sources are as follows:

- EXP and IMP are Turkey’s agricultural export and import values in terms of US$. The data are taken from TurkStat (Turkish Statistical Institute).
- \(Y_{FR}\) is the real income of trading partner \(i\) measured as GDP (constant 2000 US$). \(Y_{TR}\) is the real income of Turkey measured as GDP (constant 2000 US$). The data are collected from World Bank (2007).
- EXR is the real bilateral exchange rate between Turkish lira and trading partner’s currency. It is defined as \((P_{jr} \cdot NEX_j) / P_{TR}\), where \(P_{jr}\) is the trading partner’s Consumer Price Index (CPI), \(P_{TR}\) is Turkey’s CPI, and \(NEX_j\) is the bilateral nominal exchange rate as the number of Turkish lira per unit of trading partner’s currency. An increase in the real bilateral exchange rate therefore represents a real depreciation of Turkish lira. The data on CPIs and the nominal exchange rates are obtained from International Financial Statistics and Electronic Data Distribution System of Central Bank of the Republic of Turkey, respectively.

4. Empirical Findings

Before discussing empirical findings, it is important to note that (i) the SBC is used to determine optimal lag lengths, (ii) due to small sample size only \(k= 1, 2, ..., 4\) are considered in both unit root and cointegration analysis, and (iii) asymptotic values of \(E(Z_k)\) and \(Var(Z_k)\) are used provided by Breitung (2005).

Results of the panel unit root tests are represented in Table 1. Although the tests do not provide us a clear conclusion for which the variables are stationary in level form, all the variables are stationary in first-difference form. After justifying the variables are non-stationary in levels and integrated at most \(I(1)\), we can now move to the panel cointegration analysis.

\(^1\) Although our analysis in fact should include Spain as a major trading partner, we had to exclude it from the panel because we can not collect corresponding data, especially the bilateral exchange rate.
Table 1: Results for panel unit root tests

<table>
<thead>
<tr>
<th></th>
<th>LLC</th>
<th></th>
<th>IPS</th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Individual Fixed Effects</td>
<td>Individual Fixed Effects and Trend</td>
<td>Individual Fixed Effects</td>
<td>Individual Fixed Effects and Trend</td>
</tr>
<tr>
<td>lnEXP</td>
<td>-1.70 [0.0044]</td>
<td>-3.39 [0.0003]</td>
<td>-1.38 [0.0384]</td>
<td>-4.28 [0.0000]</td>
</tr>
<tr>
<td>lnIMP</td>
<td>-0.71 [0.2360]</td>
<td>-6.37 [0.0000]</td>
<td>0.46 [0.6773]</td>
<td>-5.79 [0.0000]</td>
</tr>
<tr>
<td>lnY_TR</td>
<td>-0.70 [0.2419]</td>
<td>0.38 [0.6492]</td>
<td>2.74 [0.9970]</td>
<td>-1.97 [0.0240]</td>
</tr>
<tr>
<td>lnY_FR</td>
<td>4.83 [0.0000]</td>
<td>-1.91 [0.0275]</td>
<td>-1.00 [0.1582]</td>
<td>-1.90 [0.0282]</td>
</tr>
<tr>
<td>lnEXR</td>
<td>-2.88 [0.0019]</td>
<td>-1.24 [0.1072]</td>
<td>-2.73 [0.0031]</td>
<td>0.03 [0.5136]</td>
</tr>
<tr>
<td>ΔlnEXP</td>
<td>-17.24 [0.0000]</td>
<td>-15.76 [0.0000]</td>
<td>-17.55 [0.0000]</td>
<td>-16.82 [0.0000]</td>
</tr>
<tr>
<td>ΔlnIMP</td>
<td>-18.68 [0.0000]</td>
<td>-17.02 [0.0000]</td>
<td>-18.90 [0.0000]</td>
<td>-18.15 [0.0000]</td>
</tr>
<tr>
<td>ΔlnY_TR</td>
<td>-13.34 [0.0000]</td>
<td>-11.72 [0.0000]</td>
<td>-18.92 [0.0000]</td>
<td>-12.82 [0.0000]</td>
</tr>
<tr>
<td>ΔlnY_FR</td>
<td>-8.88 [0.0000]</td>
<td>-8.46 [0.0000]</td>
<td>-8.40 [0.0000]</td>
<td>-7.98 [0.0000]</td>
</tr>
<tr>
<td>ΔlnEXR</td>
<td>-16.71 [0.0000]</td>
<td>-16.58 [0.0000]</td>
<td>-16.68 [0.0000]</td>
<td>-14.94 [0.0000]</td>
</tr>
</tbody>
</table>

Δ is the first-difference operator. Numbers in brackets are \( p \)-values.

Results for the panel cointegration analysis are presented in Table 2. As it was mentioned in section 2, the Pantula principle is followed to determine appropriate restrictions on deterministic terms. From Table 2, we gather that in the first row all panel rank test statistics are greater than the critical value at 10 percent level of significance and the null of zero rank is rejected for all models. As far as the second raw is considered, we observe that the first time that the null hypothesis \( r = 1 \) is not rejected for the second model. From the Pantula principle we conclude that the cointegration rank is equal to one in the panel, and the appropriate model should include unrestricted intercept and no trend for both export and import models.

Table 2: Results for the likelihood-based panel cointegration test

<table>
<thead>
<tr>
<th>Rank</th>
<th>Export</th>
<th>Import</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Model 1</td>
<td>Model 2</td>
</tr>
<tr>
<td>0</td>
<td>5.47</td>
<td>1.86</td>
</tr>
<tr>
<td>1</td>
<td>2.17</td>
<td>0.29*</td>
</tr>
<tr>
<td>2</td>
<td>1.04</td>
<td>1.34</td>
</tr>
</tbody>
</table>

Model 1: restricted intercept, no trend; Model 2: unrestricted intercept, no trend; Model 3: unrestricted intercept, restricted trend. The 1% critical value is 2.33, the 5% critical value is 1.64, and the 10% critical value is 1.28. * denotes the first time the null hypothesis is not rejected at the 10% significance level.
The estimated long-run elasticities are reported in Table 3. For the export functions, we observe that the income elasticities are positive except Italy and significant in the cases of Belgium, France, Germany, and Netherlands at 10 percent level of significance. This result implies that an increase in income level of the selected trading partners with the exception of Italy and United Kingdom is likely to increase demand for Turkish agricultural products. The long-run exchange rate elasticities are statistically significant only in the cases of France and Netherlands. This finding implies that the exchange rate is not a major factor effecting Turkey’s agricultural exports to most of her major trading partners in the long-run.

For the import functions, we observe that the income elasticities are positive in all cases and statistically significant except United Kingdom, implying that as Turkey’s income increases agricultural imports from the selected partners are likely to increase. Results for the exchange rate show that the long-run exchange rate elasticities are positive in all cases. However, the parameters are statistically significant in the cases of France, Germany, and Netherlands, implying that a depreciation of Turkish lira will increase Turkey’s demand for agricultural import from these countries.

<table>
<thead>
<tr>
<th>Countries</th>
<th>Export</th>
<th>Import</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Lag</td>
<td>$Y_{FR}$</td>
</tr>
<tr>
<td>Belgium</td>
<td>1</td>
<td>1.60 (3.65)</td>
</tr>
<tr>
<td>France</td>
<td>1</td>
<td>1.13 (3.32)</td>
</tr>
<tr>
<td>Germany</td>
<td>1</td>
<td>1.00 (2.59)</td>
</tr>
<tr>
<td>Italy</td>
<td>1</td>
<td>-4.26 (1.44)</td>
</tr>
<tr>
<td>Netherlands</td>
<td>1</td>
<td>1.77 (6.98)</td>
</tr>
<tr>
<td>United Kingdom</td>
<td>1</td>
<td>0.35 (0.64)</td>
</tr>
</tbody>
</table>

Long-run parameters are normalized with respect to export and import values. Numbers in parenthesis are absolute values of $t$-ratios.

Finally, we performed diagnostic tests for normality and autocorrelation and reported results in Table 4. It is clear that autocorrelation does not pose any problem in the estimation of all the export and import functions. The same is the case for normality except for Italy where the null of normality is rejected.
Table 4: Results for diagnostic tests

<table>
<thead>
<tr>
<th>Countries</th>
<th>Normality</th>
<th>Autocorrelation</th>
<th>Normality</th>
<th>Autocorrelation</th>
</tr>
</thead>
<tbody>
<tr>
<td>Belgium</td>
<td>9.33 [0.1558]</td>
<td>5.41 [0.7960]</td>
<td>8.41 [0.2094]</td>
<td>9.22 [0.4169]</td>
</tr>
<tr>
<td>France</td>
<td>8.90 [0.1789]</td>
<td>19.12 [0.0241]</td>
<td>13.75 [0.0325]</td>
<td>9.57 [0.3859]</td>
</tr>
<tr>
<td>Germany</td>
<td>8.16 [0.2264]</td>
<td>9.54 [0.3882]</td>
<td>9.78 [0.1340]</td>
<td>9.77 [0.3668]</td>
</tr>
<tr>
<td>Italy</td>
<td>190.88 [0.0000]</td>
<td>9.01 [0.4355]</td>
<td>190.28 [0.0000]</td>
<td>6.00 [0.3695]</td>
</tr>
<tr>
<td>Netherlands</td>
<td>9.72 [0.1366]</td>
<td>11.04 [0.2725]</td>
<td>8.59 [0.1975]</td>
<td>8.32 [0.5020]</td>
</tr>
<tr>
<td>UK</td>
<td>7.44 [0.2814]</td>
<td>14.97 [0.0916]</td>
<td>5.94 [0.4296]</td>
<td>5.76 [0.7627]</td>
</tr>
</tbody>
</table>

Numbers in brackets are p-values.

5. Conclusions

This study analyzed the exchange rate sensitivity of Turkish agricultural exports and imports on bilateral basis between Turkey and her six major trading partners in the EU for the period 1969-2005 using the likelihood-based panel cointegration analysis recently developed by Larsson and Lyhagen (1999) and Larsson et al. (2001). Results show that the exchange rate is not a major determinant of Turkish agricultural trade flows with these countries while income variables have significant impacts.

Due to the fact that trade balance is an indicator of trade performance, an implication of our results is based on international trade theory which states that devaluation improves Turkish agricultural trade balance if Marshall-Lerner condition holds. If a depreciation of Turkish lira plays a role in improving Turkish agricultural trade balance, the exchange rate elasticities should be positive in export demand functions, while they should be negative in the import demand functions. Accordingly, the ML condition does not hold in Turkish agricultural trade balance with the selected countries. There are two explanations why depreciation of the national currency does not always improve trade balance (Irandoust et al., 2006). First explanation is the demand side behavior which asserts that if demand is inelastic, a change in the exchange rate does not have the expected effect on the demand for foreign goods. Second explanation is the supply side behavior which postulates that the exporters adopt so-called “pricing-to-market-behavior” in times when exchange rate movements could damage their competitiveness. These theoretical explanations imply that Turkish policy makers should concentrate on the pricing to market phenomenon which could be a critical factor in explaining Turkish agricultural trade flows by reconsidering the role of
exchange rate policy in agricultural trade flows within the context of market structure of traded goods and pricing behavior of exporters.

From the studies investigating dynamic interrelationships between exchange rate and agricultural variables, there is a consensus of which the exchange rate is an exogenous variable for agricultural sector. Therefore, our result imply that Turkey’s macroeconomic and exchange rate policies which lead to depreciation of Turkish lira against currencies of the selected partners would deteriorate Turkish agricultural trade balance with most of these trading partners. So, Turkish agricultural trade policies should be designed by taking into account impacts of macroeconomic environment on agricultural trade flows.

It is important to note that implications suggested here have some limitations. Firstly, the empirical analysis focused on six important trading partners, which implies that these implications can not be generalized for Turkish agricultural flows. In order to better understand the role exchange rate on Turkish agricultural trade flows, future studies should use bilateral trade flows of agricultural sector as well as agricultural commodities which account for important share of Turkish agricultural trade flows by including more of major trading partners. Secondly, the models estimated here theoretically use only income and exchange rate variables as explanatory variables. However, importance of other factors such as climatic differences between Turkey and her trading partners, agricultural arable land which catches the size of agricultural production capacity, Turkish people living in trading partners, and trade agreements should be analyzed within the context of different trade flows’ models. For instance, future studies should analyze effects of set of variables as well as income and exchange rate by estimating gravity model. One can argue that importance of some factors effecting Turkish agricultural trade flows had been already analyzed by Danzinger et al. (2005), by Atici and Guloglu (2006), and by Erdem and Nazlioglu (2008). However, it is clear from this literature on Turkish agricultural trade; numbers of the studies are very limited to obtain general conclusions.
References


International Financial Statistics, IFS Online CD-ROM


