Food Price Convergence and Trade in the Balkans

by

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1. Introduction

Early theoretical economists who initiated work on purchasing power parity (PPP) believed that "once converted to a common currency, national price levels should be equal" (Rogoff, 1996). Today it is evident that such proposition does not hold for a broad range of goods. During the past few decades, investigation on why this proposition does not hold became an important topic in the theoretical and empirical macroeconomic literature. The literature developed tests that predict whether the prices will converge in the future, and if they do, how fast the convergence occurs. The literature also provides explanations for the different rate of convergence to the same price level that countries experience. The major explanations for slow convergence are summarized by Rogoff (1996), Caves, Frenkel and Jones (1990), Krugman and Obstfeld (1991), and Engel (1992). The most obvious reason for slow convergence is the presence of trade barriers, including transportation costs, tariffs, and non-tariff barriers (such as language or cultural differences). In addition, different preferences of the consumers across countries and the presence of non-traded goods in consumer price indices that are being examined may also cause slow convergence of prices.

The literature has given special attention to price convergence within the members of the European Monetary Union, as the adoption of the Euro provides a natural experiment for testing whether the prices converge faster under the single currency. However, there is still no work that focuses on price convergence in the Balkans. Yet, it is important to examine how prices move across the region for several reasons. First, if the prices move identically, this is an indication that the region has nearly exhausted its trade potentials. This follows from the fact that the law of one price is enforced through arbitrage. Second, if the prices do not move identically and converge at a slow rate, one could examine hypotheses on why that is the case. This evidence can be used to better understand the widely held notion that these countries vastly underutilize their trade potential, a notion that is so far based mainly on anecdotal evidence.[1]

As food consumed in the region is highly homogeneous, we chose to investigate the movements of food prices in order to measure convergence. The homogeneity of food in the region applies to raw food as well as to processed products, i.e. dishes prepared for home consumption. Raw food is expected to be homogeneous almost everywhere, but its composition in the final product will rarely be as identical in any other region of the world as it is in the Balkans. This is the result of the population having identical preferences for food developed over centuries of living together under same socio economic conditions.

The analysis considers only price indices and it measures only relative PPP convergence, which “requires only that the rate of growth in the exchange rate offset the differential between the rate of growth in home and foreign price indices” (Rogoff, 1996). This measure allows for a wedge between price levels across countries in the base year. Such a wedge can be explained by various factors that influence absolute prices. Different tax rates or different labor costs in food retailing and agriculture are just some of the factors that can create differences in absolute prices.

Considering food as a highly tradable good, along with population preferences described above, one would expect price indices of food to move identically across the region. When all factors such as exchange rate differences, differential VAT or sales tax rates, and tariffs are accounted for, food prices should move identically across the region if relative PPP holds. Any difference in the movement of the food prices in the neighboring countries would open profit opportunities, and arbitrage will work to balance out the differences. Therefore, perfect convergence of food prices will provide evidence of arbitrage enforcing the law of one price. Conversely, less than perfect convergence will provide evidence of potential arbitrage and trading opportunities that are not yet realized.

Our results show evidence of slow convergence in food price indices, implying that the law of one price does not hold. We evaluate the evidence of slow convergence in the light of the standard explanations suggested by the literature. In addition to the common treatment applied in the literature, we investigate whether prices in countries that are culturally similar converge faster. After considering alternative explanations for slow convergence in the food prices, we are able to conclude that the main reason for such occurrence are non-tariff barriers. Before we proceed with presenting our detailed results and offer explanation for our findings, we first present the data in following section. The econometric model is presented in Section 3. Results are presented in Section 4, and Section 5 concludes. A detailed econometric approach and summary trade statistics of the region are left for the appendix.

2. Data

The analysis focuses on the following countries: Albania, Bulgaria, Croatia, Greece, Hungary, FYR Macedonia, Romania, Serbia and Montenegro, Slovenia, and Turkey. It was desirable to include Bosnia and Herzegovina, but the unavailability of data made it nearly impossible. Monthly data on food inflation for all countries are obtained from the United Nations' *Monthly Bulletin of Statistics*. For Slovenia, the data cover only urban areas, and for Turkey the data include prices of tobacco products. Although the UN reports data for

Serbia and Montenegro, these data essentially cover Serbia only, which may cause some measurement error. However, as Montenegro accounts for only about 7% of the population of country, no serious errors are expected. The data on exchange rates were obtained from the International Monetary Fund's *International Financial Statistics* for all countries except for Serbia and Montenegro where the data come from the National Bank of Serbia.[2, 3]

The first step in our analysis is to adjust the inflation rate for changes in the exchange rate. By doing that, we essentially measure how much prices have risen (or fallen) relative to a benchmark currency. We chose to investigate inflation movements relative to the United States Dollar (USD), although using the Euro was more appealing. However, as the Euro is widely used for transactions in these countries there may be instances where the exchange-rate-adjusted inflation arises purely from movements of the Euro. Therefore, to measure the purchasing power of the domestic currency we look at the difference of how much the USD can buy in the current period relative to how much the USD could buy in the previous period.

The variable of interest then becomes:

(1) $$q_{i,t} = p_{i,t} - \xi_{i,t}$$

where:

$q_{i,t}$ - inflation rate from previous the period net of currency depreciation;

$p_{i,t}$ - inflation rate from previous the period in nominal terms;

$\xi_{i,t}$ - depreciation rate of the domestic currency against the USD from the previous period.

Figure 1 presents the deviations of inflation from the cross-sectional mean for selected countries for the sample period. The maximum and minimum deviations from cross sectional mean for each year and their differentials are presented in Table 1.

Figure 1 about here (title: *Deviation from the Cross Sectional Mean for Selected Countries)*

Table 1 about here (title: *Deviations from the Mean*)

3. Econometric Model

This section presents the model we employ to tests whether inflation indices in the sample countries follow a unit-root process, i.e. whether the inflation indices contain a stochastic trend, causing them to diverge from one another. The alternative to this hypothesis is that the indices converge after a certain period. Given the low power of the univariate unit root tests (i.e. their inability to reject the unit-root when it is false) the literature suggests using panel unit root tests developed by Im, Pesaran, and Shin (1997), denoted IPS test, and Levin and Lin (1993), denoted LL test. Their exact procedure is left for the appendix. Both tests are based on the following estimation equation suggested by Hsiao (2004):

(2) $$\Delta q_{i,t} = \alpha_i + \theta_t + \beta_i q_{i,t-1} + \sum_{j=1}^{k_i} \gamma_{i,j} \Delta q_{i,t-j} + \varepsilon_{i,t}$$

where $q_{i,t}$ is defined by equation (1). The country specific variation is captured by α_i, while θ_t captures time specific variation. The error term $\varepsilon_{i,t}$ is assumed to be normally distributed with mean 0 and variance $\sigma^2_{i,t}$. The term $\gamma_{i,j}\Delta q_{\,i,t-j}$ measures the impact of changes in q from j periods ago on the change in q in this period. Both tests constrain $\sum_{j=1}^{k_i}\gamma_{i,j} = \beta_i + 1$.

The null hypothesis is the same for both tests:

$$H_0 : \beta_i = 0 \;\forall\, i$$

If the null hypothesis is true, there is evidence that there is no convergence in the panel. If β_i is zero, it follows that the changes in the inflation rate in the current period will depend on the country specific rate of change α_i, the time effect θ_t, and the weighted average of k_i previous changes in the inflation rate.[4] In addition, there is a Gaussian shock that comes from the error term that is distributed as $N\left(0, \sigma^2_{i,t}\right)$. Consequently, Δq in this period depends only on the previous Δq and the random shock. Since all previous Δq also depend on the same variables, they follow a random walk as well. Thus, if β is 0, this means that all changes in inflation will follow a random walk and the panel will not converge as price indices in the various countries follow their own random walk.

The alternative hypothesis is different for the two tests:

$$\text{LL: } H_A : \beta_i = \beta < 0$$

$$\text{IPS: } H_A : \beta_i < 0 \textit{ for some i.}$$

The alternative hypotheses differ because the IPS test allows for differences in the coefficient estimates across countries. The IPS test then calculates the panel estimate as a cross-sectional average. The LL test, however, allows for an overall panel estimate of the coefficients. The IPS has one unique feature, which makes it less preferred to the LL test.

Namely, the IPS test requires that only one country needs to converge in order for the entire panel to converge. That is, if all but one β_i are zero and only one β_i is less than zero, the average of all β_i computed as ${}^{1}\!/_{N}\sum_{i=1}^{N}\beta_i$ will be less than zero.

The variable of interest from both tests is ρ, which is defined as:

(3) $$\rho_i \equiv \sum_{j=1}^{k_i}\gamma_{i,j} \equiv \beta_i + 1$$

The half life of a shock, which measures the speed of convergence if convergence exists, is computed as $\ln\left(\frac{1}{2}\right)\Big/\ln\left(\rho_i\right)$. The computation of the adjusted ρ is left for the appendix. The estimates of the adjusted ρ and adjusted half-life are reported along with other results in the following section, although they may not be significant. Namely, the estimates of the adjusted ρ are computed for a large number of cross sections (ideally as $N \to \infty$), and adjusting the estimated ρ turns out not to be beneficial given the cross section of only 10 countries.

4. Results

Table 2 presents the results obtained from the procedure outlined in Section 3 employing parametric bootstrap to determine the statistical significance of the coefficients, as explained in the appendix.

Table 2 about here (title: *Panel Unit Root Test Results*)

The negative and significant values of the β coefficients under both tests suggest that the countries should converge in the long run. However, the resulting estimates of the half-life of a shock are high under both tests. This presents evidence of slow convergence. Although the LL test may overestimate the half-life of a shock compared to the estimate from the IPS test, we base our arguments mainly on the results from the LL test. Namely, the *p*-values of each test in Table 2 show that the LL test is significant at nearly 0%, while the IPS test is significant only at 6.26%. In addition, we prefer using the LL test as it provides a panel estimate of β, while the IPS test provides only country specific estimates of β_i and computes a panel estimate as the average over all countries. As noted in the previous section, under the IPS test it is sufficient for only one β_i to be negative for the entire panel to converge. Underestimated half-life of a shock may therefore result from only a portion of the countries converging faster. From Table 2 under LL test we see that it takes approximately nine and a half months for a shock in price index movement in any country to diminish to half of its original magnitude. Next, we examine different theoretical and empirical approaches proposed by the literature in an attempt to explain the less than perfect adjustment of price indices.

Distance

Papell and Theodoridis (2001) find that the “distance between the countries … [is] the most important determinant” of how fast PPP convergence occurs. In the study they use the distance as a proxy for the transportation cost. If countries are far from each other, transportation costs will be high. This would in turn cause lower exchange of goods, and prices would adjust less than perfectly. However, for the many of the countries considered in this paper, the transportation costs are identical if the trade occurs within or across borders. These are all small countries and the distance to the border is never too great. In addition, some regions are closer to places that grow or distribute food in the neighboring country than to comparable places in their own country. Therefore, distance is not expected to have an impact on the speed of convergence. To verify this, for each country we performed tests repeating the exact procedure from Section 3 but including only neighboring countries. As the results were not changed significantly, we chose not to report them, noting only that the consideration of distance did not increase speed of convergence considerably.

Nontraded Goods

When studying PPP convergence it is of crucial importance to consider whether a significant portion of the price index is made up of the prices of goods that cannot be traded. Using data from US cities Cecchetti et. al. (2002) find that the portion of nontraded goods in the set is inversely related to price convergence, i.e. the higher the portion of nontraded goods in the set, the slower the convergence is. However, the impact of nontraded goods is important only when studying much broader sets of goods. In our case, food items as most other commodities are highly tradable.[5] Therefore, such considerations cannot have major implications for the price index considered here. The high tradability of food is the main reason why the prices of food converge faster than the prices of other goods in many countries. For example, Chaudhuri and Sheen (2004) found that food prices converge faster than the prices of other goods in Australia.

Socio-Economic Factors

Given the cultural differences of the region and the different political and socioeconomic ties that the countries developed with one another during the past several decades, we also test whether countries that are culturally or politically similar to one another converge faster. To do that, we group countries by religion, noting that religion is a good measure of how close countries are politically and culturally.

Although the countries are very diverse religiously, they were grouped according to the majority religion within a country. Groupings are as follows. The *Muslim Only* sub-sample contains Albania and Turkey; The *Orthodox Only* sub-sample contains Bulgaria, Greece, Macedonia, Romania, and Serbia; and The *Catholic Only* sub-sample contains Croatia, Hungary, and Slovenia. Table 3 summarizes the results obtained when each specific sub-sample was considered.

Table 3 about here (title: *Timeframe Matters; Religion Does Not)*

When it comes to religion, two tests provide different results: the LL tends to overestimate the speed of convergence, while the IPS tends to make the results more dubious by increasing the half life of a shock even more, implying even slower convergence.[6] Considering the *p*-values, we can see that the LL test again provides more significant results. It should also be noted that at least partially the faster convergence results can be attributed to smaller panel size which itself causes less variation in price indices. Nevertheless, the half life of a shock still takes on a significant value under each test and for most sub-samples. Overall, the results indicate that even after accounting for cultural differences the convergence is still slow.

We also consider the possibility that increased integration in recent years will have an effect. This applies to the integration of the region toward the European Union (EU), but also to integration of the region itself. It has been a common practice of EU officials to emphasize the importance of integration of the region itself, and various EU institutions usually provide incentives for countries of the region to sign bilateral or multilateral free trade agreements. Although such policies of the EU continue even to this date, EU officials seem fairly enthusiastic in noting that in terms of free trade "since … 2001 a lot has been accomplished" (Stability Pact Trade Working Group, 2003). If this is true, we would expect faster convergence in more recent years. Slow convergence is expected to come from earlier periods when the countries were less integrated. We proceed by testing this possibility by grouping the panel into *Early Period* and *Later Period*. The *Early Period* includes observations from November 1999 to December 2001. The *Later Period* includes observations from January 2002 to December 2004. Table 3 indicates that there is evidence of faster convergence when compared to the original time span. However, the half life of a shock still takes on a significant value, indicating that convergence is less than perfect.

We should note that faster convergence in the two sub-samples can be entirely due to the sub-sample size: both sub-samples have considerably fewer observations when compared to original sample size. Limiting the number of observations over time to a fairly small number may itself cause convergence to be faster. Namely, in such a short timeframe not all shocks are accounted for, and estimates of half life will be greatly underestimated. Table 3 also indicates that there is evidence of faster convergence in the *Early Period* when compared to the *Later Period* when either test is employed. This in turn indicates that overall the slow convergence comes mostly from the later period. A priori, this was not expected, as the region started becoming more integrated from 2001 onward.

In addition to the factors we considered above, there are many more that could potentially influence the speed of convergence of the price indices, e.g. the different VAT, sales tax, or tariff rates in these countries. These factors, however, should not influence the speed of convergence as long as the specific tax or tariff rate is not changing over time (Cecchetti et al., 2002). The argument follows from the fact that the analysis focuses on price level changes, not absolute price changes – any effect that different tax or tariff policies have on absolute prices will be offset when changes in price indices are considered. Absent other factors, relative prices should move the same regardless of tax or tariff rates.

Non-tariff barriers to trade are the only remaining factors that can be crucial in determining the volume of trade and the speed of convergence of price indices. Such barriers to trade are interesting not only because they cannot be explicitly measured, but also because they place no explicit cost on potential trades. However, non-tariff barriers increase the cost of arbitrage substantially, and lower the incentive to trade by lowering profit opportunities (Faber and Stokman, 2004). Based on the analysis we performed, when all other factors are accounted for it appears that the only remaining factors that can potentially explain slow convergence in price indices are non-tariff barriers to trade. Among non-tariff barriers, incomplete information and bureaucratic difficulties probably have the highest negative impact on trade incentives. Incomplete information refers to traders in the region being unaware of the trade potentials that can be realized in the countries of the region. This is perhaps the result of importers and exporters focusing mainly on the EU countries under the belief that the greatest profit opportunities are there. Such belief may drop a shadow on the potentials that wait to be realized within the region itself. What is more important is that such belief continues to hold over time. Our results show evidence of slower convergence starting with 2002 as compared to the pre-2002 period, implying that the activity of importers and exporters may be shifting outside the region over time.

Bureaucratic difficulties are a very broad category: they can include everything from large amount of forms that need to be filed in order for goods to cross the border, to different and difficult to understand local product requirements. The presence of such barriers can cause significant deviations from the law of one price and may confirm the common belief that these countries vastly underutilize their trade potential.

5. Conclusion

The Balkans remains the region with the lowest prosperity and highest security risks in Europe. One possibility of increasing prosperity and lowering the security risks in the region is trade. Theory and empirical work suggest that countries more open to trade have higher prosperity, if we measure it by economic growth. Trade also makes countries more integrated and more dependent on one another. This dependence increases the cost of any potential conflict, and lowers security risks. Despite many free trade negotiations among the countries of the region and continuous policies that aim to liberalize trade, the general perception is that the countries vastly underutilize their trade potentials. Low trade volume can be either attributed to lack of trading opportunities, or to high degree of barriers to trade. Using PPP convergence as a measure, we show that there are still trade opportunities to be realized. Therefore, the low trade volume together with existing trading opportunities indicate the presence of barriers to trade that prevent these trade opportunities to be realized.

Although there is no direct policy implication that follows from our results, it should be noted that simply promising to liberalize trade may not lead to full utilization of trade potential. Governments in the region may consider taking a more active role in promoting trade. Namely, governments can present people with different opportunities that could be realized in the region. At the same time, governments of the region should commit to lower bureaucratic difficulties. Such policies are already known to exist between countries of the region and the countries already in the EU. They therefore may serve as a good reference point on how government policies may encourage trade inside the region.

Appendix 1: Econometrics

A priori choices of k_i are made according to procedure suggested by Campbell and Perron (1991). Then, optimal k_i are obtained in the following manner: if k_i th lag of $\Delta q_{i,t}$ has significant impact on $\Delta q_{i,t}$, k_i is chosen as optimal; if it does not have significant impact, k_i is reduced by 1. Procedure is repeated until last lagged value of $\Delta q_{i,t}$ appears significant. Values of k_i are allowed to differ across countries. After k_i are obtained, we first estimate equation (2). Then, parametric bootstrap for the LL and IPS tests and their significance follows procedure suggested by Cecchetti et al. (2002) and Hsiao (2004).

Under the null hypothesis, data generating process is assumed to be:

(A.1) $$\Delta q_{i,t} = \mu_i + \sum_{j=1}^{k_i} \Delta q_{i,t-j} + \varepsilon_{i,t}$$

Following simple ordinary least square (OLS) procedure of equation (2), variance-covariance matrix Σ is obtained from residuals as $\hat{\Sigma} = N\left(\varepsilon_t^{'}\varepsilon_t\right)$. Then, $T+100$ innovations are drawn from multivariate normal distribution as $\tilde{\varepsilon}_{i,t} \sim N\left(0, \hat{\Sigma}\right)$. Pseudo observations $\Delta\hat{q}_{i,t}$ and $\hat{q}_{i,t}$ are generated according to (A.1). First 100 pseudo observations are dropped, and then LL and IPS tests are performed on the remaining data to obtain bootstrap distribution.

Levin and Lin Test

Levin and Lin (1993) suggest different approaches in testing the panel data for convergence, but we follow the procedure outlined by Cecchetti et al. (2002):

1. Subtract cross sectional mean from each observation to remove common time effect:

(A.2) $$\tilde{q}_{i,t} = q_{i,t} - \tfrac{1}{N}\sum_{i=1}^{N} q_{i,t}$$

2. Now, for each country:

    a. Regress $\Delta\tilde{q}_{i,t}$ on a constant and k_i lagged values of $\Delta\tilde{q}_{i,t}$ and obtain residuals denoted $\hat{e}_{i,t}$;

    b. Regress $\tilde{q}_{i,t-1}$ on a constant and k_i lagged values of $\Delta\tilde{q}_{i,t}$ and obtain residuals denoted $\hat{v}_{i,t-1}$;

    c. Regress (with no constant) $\hat{e}_{i,t}$ on $\hat{v}_{i,t-1}$, and obtain residuals denoted $\hat{\omega}_{i,t}$.

    Compute standard error of the regression as $\hat{\sigma}_{ei} = \sqrt{\left(T-k_i-1\right)^{-1}\sum_{t=k_i+2}^{T}\hat{\omega}_{i,t}^2}$.

Normalize $\hat{e}_{i,t}$ and $\hat{v}_{i,t-1}$, and denote normalized values as $\tilde{e}_{i,t} = \hat{e}_{i,t} / \hat{\sigma}_{ei}$ and $\tilde{v}_{i,t} = \hat{v}_{i,t} / \hat{\sigma}_{ei}$.

3. Run panel OLS regression as follows:

(A.3) $$\tilde{e}_{i,t} = \beta \tilde{v}_{i,t-1} + u_{i,t}$$

Reported *t*-statistic of β coefficient is the τ for LL test.

Im, Pesaran, and Shin Test

To conduct test proposed by Im, Pesaran, and Shin (1997), time trend is removed as in equation (A.2). Then, for each country, we run augmented Dickey-Fuller regression of $\Delta\tilde{q}_{i,t}$ on a constant, $\tilde{q}_{i,t-1}$, k_i lagged values of $\Delta\tilde{q}_{i,t}$. Then, t_i is obtained as *t*-statistic from the coefficient of $\tilde{q}_{i,t-1}$. Overall $\overline{t}$ is obtained as $\overline{t} = \frac{1}{N}\sum_{i=1}^{N} t_i$.

Procedure outlined above is repeated 2000 times. Realizations of τ and $\overline{t}$ are used to calculate the precision of the tests, and their corresponding coefficients are used to calculate estimated half-lives of shocks.

Computation of Adjusted ρ

Adjusted $\hat{\rho}$ and adjusted $\hat{\hat{\rho}}$ are calculated using formula suggested by Nickell (1981). We calculate the difference between true and estimated ρ and then adjust our estimates. Procedure is outlined by the following equation:

$$\operatorname{plim}_{N\to\infty}\left(\hat{\rho}-\rho\right)=\frac{\dfrac{-(1+\rho)}{(T-1)}\left(1-\dfrac{1}{T}\dfrac{\left(1-\rho^{T}\right)}{(1-\rho)}\right)}{1-\dfrac{\dfrac{2\rho}{T}\dfrac{\left(1-\rho^{T}\right)}{(1-\rho)}}{(1-\rho)(T-1)}}$$

Appendix 2: Trade Volume

Looking at the general patterns of exports and imports that prevail in the region, it is not hard to conclude that the countries are not important trading partners to one another. Table 4 summarizes importance of trade (in terms of rank and percentage of total imports of exports) among the countries in the region. With the exception of FYR Macedonia (whose first exporting partner is Serbia and Montenegro and first importing partner is Greece) and Albania (whose second importing partner is Greece), we see that countries are not important to one another when it comes to trade.

Table 4 about here (title: *Importance of Imports and Exports across Countries in 2004*)

Notes

1. For reference, Appendix 2 includes table that presents trade volume and relative importance of countries to one another in terms of trade.
2. Data from National bank of Serbia is compared with data from Bloomberg and OANDA Corporation for the periods prior to June 2001 when Dinar was fixed. Either method shows nearly identical results.
3. In the panel size of 620 observations, only 23 observations were missing. As unit root tests would provide inconsistent results if these time periods were dropped, we proceeded by averaging missing observations using a cross sectional average, or by assigning them the value of the average of one period before and one period ahead of the country where the data were missing. Either procedure did not change the results in meaningful ways. In addition, such treatment of the data would bias the results toward convergence.
4. This is weighted average as coefficients are constrained to sum to one.
5. Tradability of food items here refers only to ability to exchange the goods at different markets. We do not suggest there exist no barriers to exchange them.
6. Half life is again measured using ρ. Adjusted ρ and adjusted half life are computed but not reported, as they are very insignificant for sample sizes of only 2, 3, and 5.

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Figure 1: Deviation from the Cross Sectional Mean for Selected Countries

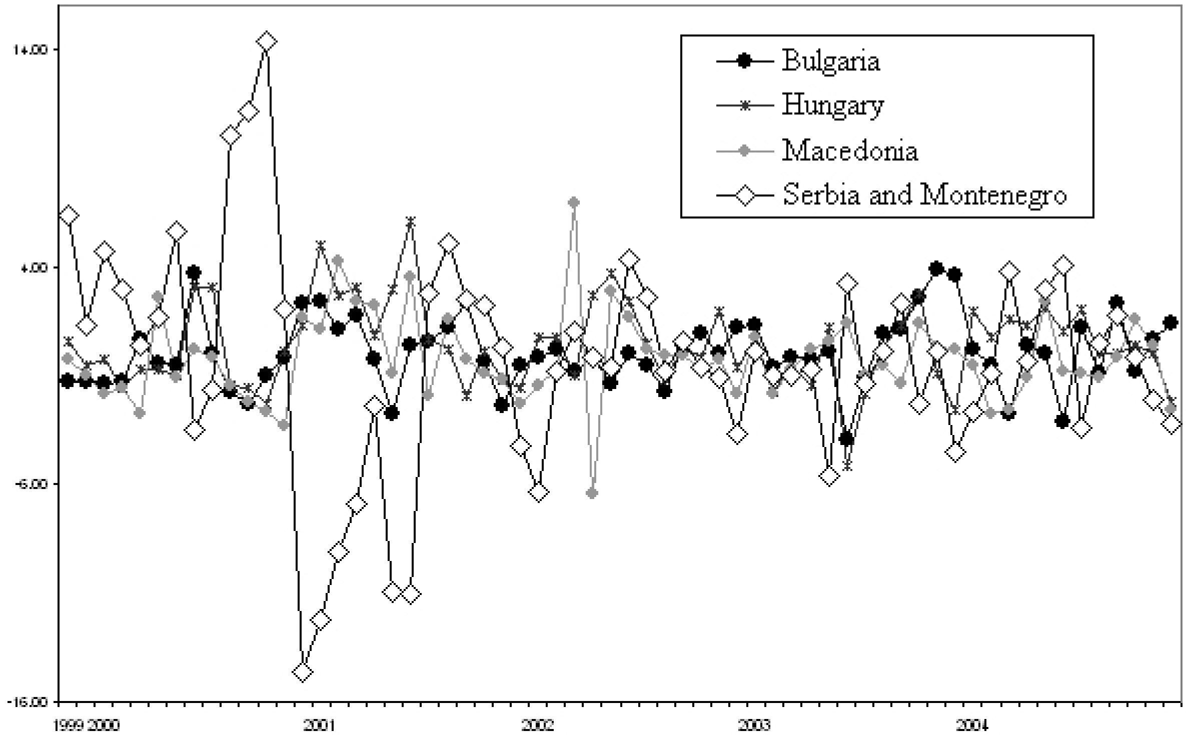

Table 1: Deviations from the Mean

Year	Maximum	Country	Minimum	Country	Differential
2000	2.74	Serbia	-1.06	Macedonia	3.80
2001	1.40	Hungary	-3.66	Serbia	5.06
2002	0.81	Hungary	-0.49	Serbia	1.31
2003	0.84	Turkey	-0.90	Serbia	1.73
2004	0.73	Hungary	-0.69	Macedonia	1.42

Note: Maximum and minimum are from country's average deviation from the cross-sectional mean during corresponding year.

Table 2: Panel Unit Root Test Results

Levin and Lin					
τ	p-value	$\hat{\rho}$	Adjusted $\hat{\rho}$	Half-life	Adjusted half-life
-4.1544	0.0000	0.9295	0.9024	9.4817	6.7478

Im, Pesaran, and Shin					
$\overline{t}$	p-value	$\hat{\bar{\rho}}$	Adjusted $\hat{\bar{\rho}}$	Half-life	Adjusted half-life
-1.5355	0.0626	0.8899	0.8624	5.9445	4.6832

Table 3: Timeframe Matters; Religion Does Not

		Muslim Only	Orthodox Only	Catholic Only	Early Period	Later Period
LL Test	Half life	6.9447	7.9494	7.7655	3.6018	4.4876
	p-value	0.0166	0.0010	0.0079	0.0000	0.0000
	% Change from original	*-26.76%*	*-16.16%*	*-18.10%*	*-62.01%*	*-52.67%*

		Muslim Only	Orthodox Only	Catholic Only	Early Period	Later Period
IPS Test	Half life	6.9447	5.5999	6.1483	2.23	2.6452
	p-value	0.0676	0.0604	0.065	0.0586	0.0653
	% Change from original	*16.83%*	*-5.80%*	*3.43%*	*-62.49%*	*-55.50%*

Note: Half life shows value obtained from specific sample, and it is followed by *p*-value of the corresponding coefficient. *% Change from original* represents percentage change form original estimate of half life based on entire panel.

Table 4: Importance of Imports and Exports across Countries in 2004

	Albania		**Bulgaria**		**Croatia**		**Greece**		**Hungary**		**FYR Macedonia**		**Romania**		**Serbia and Montenegro**		**Slovenia**		**Turkey**	
	E	*I*	*E*	*I*	*E*	*I*	*E*	*I*	*E*	*I*	*E*	*I*	*E*	*I*	*E*	*I*	*E*	*I*	*E*	*I*
Albania			-	-	-	-	-	-	-	-	-	-	-	-	-	-	-	-	-	-
Bulgaria	-	-			-	-	4 (6.3)	-	-	-	-	5 (8.1)	-	-	-	5 (4.7)	-	-	-	-
Croatia	-	-	-	-			-	-	-	-	4 (6.9)	-	-	-	-	-	5 (7.4)	-	-	-
Greece	-	2nd (19.8%)	5 (5.6)	4 (7.5)	-	-			-	-	3 (8.9)	1 (15.4)	-	-	4 (6.7)	-	-	-	-	-
Hungary	-	-	-	-	-	-	-	-			-	-	-	-	-	-	-	-	-	-
FYR Macedonia	-	-	-	-	-	-	-	-	-	-			-	-	-	-	-	-	-	-
Romania	-	-	-	-	-	-	-	-	-	-	-	7 (4.7)			-	-	-	-	-	-
Serbia and Montenegro	-	-	-	-	-	-	-	-	-	-	1 (31.4)	3 (10.4)	-	-			-	-	-	-
Slovenia	-	-	-	-	5 (7.6)	4 (7.1)	-	-	-	-	-	4 (8.6)	-	-	6 (4.1)	4 (6.7)			-	-
Turkey	-	3 (7.7)	3 (9.4)	5 (6.9)	-	-	7 (4.5)	-	-	-	-	6 (6)	4 (7)	5 (4.2)	-	-	-	-		

Notes: For each country in first row *E* presents exports to the country in the corresponding row, while *I* presents imports from the country in corresponding row. First number presents ranking, while number in parentheses presents percentage of total imports or exports of the country on top. Source: *CIA World Factbook*, 2005.