

Business cycle synchronization between the Western Balkans and the European Union

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This paper investigates the degree of business cycle synchronization (BCS) between the Western Balkans and the European Union. While the degree of synchronization had been low or even negative before 2000, business cycles have clearly converged ever since. Since 2009, in particular, BCS with the EU has been high for all Western Balkan countries. In addition, this paper empirically tests which factors were responsible for the convergence process. Using a period-by-period correlation index for the time span from 1994 to 2013, we find that foreign trade is the most important positive contributor to business cycle convergence. Although fiscal differences are associated with negative BCS in the same year, they have a positive influence in subsequent periods. In contrast, our results suggest that financial flows lead to business cycle decoupling because of their procyclical behavior in the respective domestic economy. The same relationship applies, according to our analysis, to remittances, whose impact on BCS is an underresearched topic in the empirical literature. We find that remittances from the EU to the Western Balkans behave similarly to financial flows, which supports the hypothesis that remittances are sent home to take advantage of favorable economic conditions.

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Four of the seven Western Balkans economies² follow a monetary policy that is very closely linked to that of the euro area, either through the use of the euro as their official currency (Montenegro and Kosovo) or through fixed exchange rate regimes (euro-based currency board in Bosnia and Herzegovina, euro peg in FYR Macedonia). Although Albania, Croatia and Serbia operate under a managed, or rather free-floating, exchange rate regime, respectively, they also face in practice substantial monetary policy constraints, given the high asset and liability euroization in their banking systems.

According to the optimum currency area theory (OCA) developed by Mundell (1961), McKinnon (1963) and Kenen (1969), one prerequisite for the efficient use of a common currency or for following the monetary policy of another country or currency area is that the business cycles of the countries involved are sufficiently synchronized. Otherwise – if output patterns are divergent – monetary policy cannot bring about optimal reactions for each country at the same time. When evaluating the economic costs of a country's lack

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² In the regional definition of the Western Balkans, we follow the IMF (see https://www.imf.org/external/pubs/ft/reo/2015/eur/eng/pdf/erei_sr_030915.pdf) and include Albania, Bosnia and Herzegovina, Croatia, FYR Macedonia, Kosovo, Montenegro and Serbia.

of, or tight constraint on, independent monetary policy, it is therefore important to know the degree of BCS of the respective country with the euro area (and, in a broader, forward-looking sense, with prospective euro area countries).

This paper is structured as follows. Section 1 provides an overview of the relevant literature. In section 2, we investigate the degree of BCS between the Western Balkan countries and the EU-25 aggregate to find out which countries exhibit higher or lower BSC, and whether a convergence process can be identified. Section 3 discusses the main drivers and transmission channels of BCS and the choice of explanatory variables. Subsequently, we employ a regression model for analyzing what factors drive the development of BCS over time. Finally, in section 4, we draw conclusions from our analysis.

1 Literature overview

The degree of BCS between countries is driven by two main factors: on the one hand, it depends on the presence and dominance of common transnational shocks over idiosyncratic (i.e. country-specific) shocks. On the other hand, idiosyncratic shocks by themselves can be spread through certain transmission channels across countries. Drivers of BCS are a well-researched topic in the empirical literature. Focusing mainly on industrial countries (see e.g. Frankel and Rose, 1998; Darvas et al., 2005; Artis et al., 2008; Inklaar et al., 2008), the literature finds that trade is one of the main transmission channels of BCS. The evidence for other factors like industrial specialization or fiscal and monetary policy, which have been proposed by theoretical literature and investigated empirically, is mixed (De Haan et al, 2008). Prompted by the accession of ten Central, Eastern and Southeastern (CESEE) countries to the EU in the 2010s, several studies also investigated the patterns and drivers of synchronization between the industrialized and transition countries in Europe (see e.g. Artis et al., 2008; Babetskii, 2005; Crespo Cuaresma et al., 2011). Their results suggest that the determinants of BCS between emerging markets and industrialized countries seem to be similar to those that dominate BCS between industrial countries. Distinguishing between industrial and developing countries, Calderón et al. (2007) find, however, that the impact of trade integration is higher for BCS between industrial countries than for BCS between developing countries or between “mixed” pairs. Another strand of business cycle literature is concerned with the endogeneity of OCA criteria, as pointed out first by Frankel and Rose (1998). This endogeneity implies that a country is more likely to fulfill the OCA criteria after having joined a currency union than before accession (see Gächter and Riedl, 2014, for a discussion).

Only a few studies so far have covered the degree of BCS of the Western Balkan economies with the euro area or the European Union (EU). Velickovsky (2013) investigates shock synchronization between Western Balkan countries (Albania, Croatia, Macedonia and Serbia) and the euro area by two approaches. First, he calculates correlation coefficients between the four Western Balkan countries and the euro area and finds that output correlation is highest between Croatia and the euro area and between

FYR Macedonia and the euro area whereas correlation between Albania and Serbia vis-à-vis the euro area is comparably low. Second, the author estimates time-varying coefficients of shock symmetry (both supply and demand shocks) between the Western Balkans and the euro area in a vector autoregressive (VAR) framework. Gouveia (2014) uses a data set of eight Balkan economies (including Croatia, FYR Macedonia and Serbia from the Western Balkans) and compares various measures of trade intensity and BCS between these countries and the euro area average. For Croatia and Serbia, the degree of BCS is found to be moderate and well below the average of intra-euro area correlation, while FYR Macedonia exhibits higher output synchronization with the euro area. With respect to output volatility, the author concludes that the volatility of business cycles is substantially higher in the Balkan countries than in the euro area.

To the best of our knowledge, there is no study on BCS that covers all Western Balkan countries nor one that investigates the determinants of business cycle convergence between the Western Balkans and the euro area or the EU. This paper aims to fill this gap in the empirical literature, i.e. to identify the degree of synchronization between the Western Balkans and the EU, to find how BSC has changed since the beginning of transition and to identify the main drivers of business cycle convergence.

More precisely, our analysis covers the EU-25, i.e. all EU Member States with the exception of Croatia, which joined the EU only in mid-2013 (and is covered here in the group of Western Balkan countries), and Denmark and the United Kingdom, which have been granted an opt-out clause and are thus not required to participate in the Stage Three of Economic and Monetary Union (EMU). All other EU Member States are required to join the euro area once they fulfill the convergence criteria, as 19 of them have done so far. We investigate BCS between the Western Balkans and the current euro area as well as all EU Member States (except Croatia) that are obliged to adopt the euro and thus are the relevant counterpart for the Western Balkans' future business cycle convergence and economic integration. A further advantage of this approach is that, by doing so, we obtain a larger data set, which is particularly helpful in underpinning the robustness of our results regarding the determinants of BCS.

In our analysis, we make use of some recent advances in the business cycle literature. First, we use a new time-varying correlation index developed by Cerqueira and Martins (2009) and Cerqueira (2013) to analyze the convergence process on a yearly basis. As can be seen in the results in the subsequent sections, using a time-varying correlation index considerably improves the measurement of BCS, as synchronization varies greatly over time and this variation could not be captured by a single correlation coefficient spanning the whole observation period. Additionally, we obtain a panel data set instead of a cross-sectional sample for the regressions, which allows us to explore the full-time variability of the data. Most studies that identify determinants of BCS use a two-step instrumental variables approach; however, we employ a system Generalized Method of Moments (GMM) estimator developed by Blundell and Bond (1998) as it offers several advantages. First, it allows us to draw a large number of instruments from within the data set by instrumenting endogenous variables with their own lagged values. Second, additional time-invariant instruments can be included in the regression (in contrast to the difference

GMM estimator by Arellano and Bond, 1991). Additionally, the present data set is a small T , large N panel data set, for which the estimator is well suited as it controls for the dynamic panel bias (Roodman, 2009). To identify the determinants of BSC, we start by looking at well-known factors identified in the literature (trade, specialization, fiscal policy) but extend the choice of variables by adding other potential channels that might be relevant especially for the Western Balkans. We incorporate common monetary policy and financial flows in the regressions. In addition, we include remittances as they constitute a large part of income in the Western Balkan countries, are less volatile than other financial flows to the region (Petreski and Jovanovic, 2013) and, as already argued by Barajas et al. (2012), constitute an important BCS channel. According to the World Bank, remittances amounted to more than 16% of GDP in Kosovo and to more than 10% of GDP in Bosnia and Herzegovina in 2013. The inflow of remittances was quite substantial also in Albania, Montenegro and Serbia (between about 6% to 8% of GDP). Only Croatia and FYR Macedonia registered inflows of below 4% in 2013.

2 Business cycle synchronization

2.1 Measurement issues

In the business cycle literature, a variety of indicators of economic activity – like GDP (in levels or first differences), industrial production or consumption measures – have been used to identify the cyclical component of economy activity (see Darvas and Szapáry, 2008, for a comparison and discussion). Since the Western Balkans underwent a process of de-industrialization particularly at the beginning of the transition period in the 1990s and since industrial production does not include all sectors of the economy, we do not consider industrial production a good indicator of economic activity. We therefore make use of annual GDP data (year on year), as quarterly GDP data do not provide long time series and are not even available for some of the countries covered. Instead of using growth rates, we use the logs of real GDP data in levels as these are better suited for heterogeneous samples of countries at different stages of economic development (Gächter and Riedl, 2014).

To separate the cyclical component (i.e. the output gap) from the trend component (i.e. potential output), we use the Hodrick-Prescott (HP) filter as it is the standard method and easy to implement. Moreover, the resulting cyclical components are similar to those of the band-pass filter (Belke and Heine, 2006; De Haan et al., 2008). Following the Ravn-Uhlig rule, the smoothing parameter is set to 6.25 as proposed for yearly data. One drawback of the HP filter is that it delivers suboptimal results at the end of the sample (see e.g. Mise et al., 2005). To overcome this problem, we complement the time series for each country by forecasts from the IMF World Economic Outlook until 2019. The filtered cyclical components are tested for stationarity with the augmented Dickey-Fuller unit root test including a constant. With the exception of the cyclical component for Greece, all cyclical components in the sample are stationary at a 5% confidence level.

Most studies on this issue are affected by the nonavailability of a year-by-year index for BCS, which is why they had to investigate the topic on a cross-sectional basis. To account

for at least some time variability, most studies use moving averages, sample period splits and other methods. The lack of year-by-year BCS indices can be overcome by using the period correlation index developed by Cerqueira and Martins (2009):

(I)

$$\rho_{ij,t} = 1 - \frac{1}{2} \left(\frac{d_{j,t} - \bar{d}_j}{\sqrt{\frac{1}{T} \sum_{t=1}^T (d_{j,t} - \bar{d}_j)^2}} - \frac{d_{i,t} - \bar{d}_i}{\sqrt{\frac{1}{T} \sum_{t=1}^T (d_{i,t} - \bar{d}_i)^2}} \right)^2$$

Here, d_i and d_j denote any two time series and \bar{d}_j , \bar{d}_i denote the respective averages over time. The index thus measures the correlation between d_i and d_j at each point in time ($t = 1, \dots, T$). Taking the average of $\rho_{ij,t}$ equals the linear correlation index ρ_{ij} conventionally used in cross-sectional studies. The index is of an asymmetric nature with $\max(\rho_{ij,t}) = 1$ and $\min(\rho_{ij,t}) = 3 - 2T$ (see Cerqueira, 2013).

2.2 Data sample and descriptive results

GDP data (in euro) at constant prices are extracted from the IMF World Economic Outlook Database. Our country sample comprises 25 EU Member States (EU-25, i.e. all current members excluding Croatia, Denmark and the United Kingdom) and the Western Balkan countries (Albania, Bosnia and Herzegovina, Croatia, Kosovo³, FYR Macedonia, Montenegro and Serbia). The business cycle measures are calculated for the maximum period from 1989 to 2013.⁴

Table 1 summarizes the yearly measures of correlation between the Western Balkans and the EU-25 aggregate for different subperiods. When comparing the development of this synchronization index over the four subperiods (transition period, precrisis period, crisis peak and crisis aftermath), we can indeed identify a convergence process. Except in the crisis peak subperiod, the correlation index has clearly increased since the transition period.

³ Due to data limitations, Kosovo is included in the descriptive analysis of BCS but not in the empirical part.

⁴ For most Western Balkan countries, GDP data series start later than 1989.

Table 1

Synchronization of the Western Balkans with the EU-25 aggregate in different subperiods¹

Country ²	Transition period (until 2000)	Precrisis period (2001–08)	Crisis peak (2009)	Crisis aftermath (2010–13)	Overall (first year available: 2013)
Montenegro	n/a	0.860	−0.668	0.871	0.760
Croatia	0.401	0.765	−0.778	0.909	0.572
FYR Macedonia	0.473	0.669	−1.383	0.871	0.532
Bosnia and Herzegovina	0.313	0.720	−1.615	0.832	0.526
Serbia	−1.381	0.568	−0.869	0.839	0.181
Kosovo	n/a	0.195	−3.072	0.908	0.175
Albania	−0.073	0.287	−2.424	0.850	0.096

Source: Authors' calculations.

¹ To be consistent with the empirical part of our paper, table 1 displays the aggregate of the EU-25 and not that of the euro area. However, synchronization between the Western Balkan countries and the euro area aggregate is very similar.

² Sorted by overall synchronization level.

Note: Data are available from 1989 for Albania, from 1992 for Croatia and FYR Macedonia, from 1998 for Bosnia and Herzegovina as well as Serbia, and from 2000 for Kosovo and Montenegro.

The different subperiods are characterized by specific global and country-specific shocks. During the transformation period from 1989 to 1995, the Western Balkans were marked by the break-up of former Yugoslavia, trade interruptions and the Balkan Wars. The latter also led to EU financial and economic sanctions against Montenegro and Serbia (World Bank, 2004). Also in this period, Albania started to ease its isolation policy. These country-specific shocks of the 1990s make the analysis of BCS of the Western Balkans a difficult task. After 1995, the region saw some economic and political stability but experienced transformational recession. In addition, the limited availability of GDP data for this period also puts limits on the interpretation of business cycle correlations.

From 2001 to 2008, BCS with the EU-25 increased in particular in Bosnia and Herzegovina, Croatia, FYR Macedonia and Montenegro. In this period, the EU integration process started in all Western Balkan countries, but it progressed at different speeds. Supposedly, closer ties with the EU and increasing trade relations have had a positive impact on business cycle co-movements. At the Thessaloniki European Council summit in 2003, the EU confirmed that the future of the Western Balkans lies within the EU and granted potential EU candidate status to the Western Balkan countries (European Council, 2003). For FYR Macedonia, however, an agreement on trade and trade-related matters had already entered into force in 2001, and the Stabilisation and Association

Agreement (SAA) became effective in 2004. In the remaining countries of our sample, the SAA became effective at a later stage (European Commission, 2015).⁵

For Albania, Kosovo and Serbia, synchronization with the EU-25 was much weaker and more volatile. Kosovo strongly decoupled from the EU-25 aggregate in 2006, and Albania in 2007. Kosovo showed pronounced cyclical movements over this period, posting very high GDP growth (more than 8%) in 2007, while growth was much lower in the years before 2007 and afterward. This helps explain the large fluctuations of Kosovo's BCS. Albania experienced no strong cycle movement over time. However, in 2007, growth was relatively weak in Albania compared to a booming EU-25, which is mirrored in a low co-movement of the respective business cycles.

In 2009, the economic and financial crisis hit the EU countries hard, as reflected in the slump of the cyclical component of the EU-25 aggregate and eventually in the decoupling of the business cycle of all Western Balkan countries from the EU-25. The Western Balkans also experienced some economic downturn, but it occurred later and was less pronounced than in the EU-25. According to Bonetto et al. (2009), the Western Balkan countries were partially protected from the economic and financial crisis of 2009 as they had a low exposure to international financial markets and the foreign banks active in the region were strong capitalized. Albania and Kosovo even overcame the crisis years without dipping into a recession.

From 2010 to 2013, the convergence of the business cycles of the EU-25 and the Western Balkans increased. Even the group of Western Balkan countries recording less synchronized cycles exhibited a high degree of synchronization in that period.

As discussed above, studies which also analyze the degree of BCS between the EU and the Western Balkans generally come to similar results: FYR Macedonia shows relative strong co-movements with the EU whereas EU correlations with Albania and Serbia are comparatively low. Results for Croatia are ambiguous. Furthermore, Gouveia (2014) also shows that business cycle correlation has in general increased the over time between the Western Balkan countries and the EU.

3 Determinants of business cycle synchronization

The result of our descriptive analysis, namely that the business cycles of the Western Balkans have clearly converged with the EU business cycle over the past 15 years, leads to the question which factors drove this convergence process. We therefore proceed to empirically test the determinants of BCS.

3.1 Empirical model

In order to investigate determinants of BCS between the EU and the Western Balkans, we use the following model:

⁵ So far, no SAA has been signed between the EU and Kosovo (European Commission, 2015).

(II)

$$Correl_{ij,t} = \alpha + \beta Correl_{ij,t-1} + Z'_{ij,t}\gamma + \mu_{ij} + \lambda_t + v_{ij,t}$$

where $Correl_{ij,t}$ denotes the bilateral correlation index of BCS between countries i and j in year t . $Z'_{ij,t}$ is a matrix consisting of the potential determinants of BCS such as bilateral trade, asymmetry of production, fiscal differences, common monetary policy, FDI, bank flows and remittances. Data are available yearly in an unbalanced panel for 162 country pairs with a maximum time span from 1989 to 2013. To control for autocorrelation, the one-period lagged BCS $Correl_{ij,t-1}$ is included in the model. Additionally, we control for country-pair fixed effects μ_{ij} to account for specific unobservable country-pair factors, and time effects λ_t to account for common global shocks. Our main interest lies in the signs and magnitudes of vector γ , which indicates what drives BCS between the EU and the Western Balkans. While some authors (Imbs, 2006, and Déés and Zorell, 2011) investigate the determinants of BSC with a system of equations to account for indirect effects as well, we focus on a single equation approach. However, future research on this topic might expand our approach into a multi-equation system.

Frankel and Rose (1998) pointed out that endogeneity plays an important role in the relationship between trade integration and BCS. They argue that countries with stronger trade integration and countries with similar output patterns are more likely to join a currency union, and joining a currency union in turn increases trade integration and business cycle correlation. Following this argument, several endogeneity issues with respect to BCS have been discussed in the literature, not only with respect to trade integration but also regarding financial integration (De Grauwe and Mongelli, 2005) or remittance flows (Frankel, 2011). Most studies tackle the endogeneity issue by using two-step instrumental variable approaches. However, we follow Cerqueira and Martins (2009) and make use of a two-step system GMM estimator developed by Blundell and Bond (1998), which offers several advantages. First, it allows us to draw a large number of instruments from within the data set by instrumenting endogeneous variables with their own lagged values. Second, additional time-invariant instruments can be included in the regression (in contrast to the difference GMM estimator used by Arellano and Bond (1991)).

⁶ As pointed out by Cerqueira (2013), one shortcoming of the index (I) is that it is asymmetric, which could lead to biased results when used in regressions. Thus, for the purpose of the regressions, we use the following transformation

developed by Cerqueira (2013): $\rho_{ij,t} = \frac{1}{2} \ln \left(\frac{1 + \frac{\rho_{ij,t}}{2T-3}}{1 - \rho_{ij,t}} \right)$ that yields a nonbounded index with a symmetric support

around 0 and a symmetric range between $-\infty$ and $+\infty$.

3.2 Choice of explanatory variables

In this section we briefly describe the set of variables identified in the literature as important determinants of business cycle convergence. There is a broad consensus that trade integration is an important driver of BCS. It is argued that the elimination of trade barriers results in stronger transmission of demand shocks and eventually higher business cycle co-movement (Frankel and Rose, 1998). For testing the impact of trade on BCS between the EU-25 and the Western Balkans, we use the IMF’s database “Direction of Trade Statistics” (DOTS). DOTS provides data for the calculation of bilateral trade intensities between individual countries of the EU-25 and the Western Balkans. We calculate bilateral trade intensity as follows:

(III)

$$BTI_{ij,t} = \left(\frac{Exports_{ij,t} + Imports_{ij,t}}{Trade_{i,t} + Trade_{j,t}} \right)$$

$BTI_{ij,t}$ is obtained by dividing the sum of bilateral trade flows between i and j by the sum of total trade of country i , $Trade_{i,t}$, $Trade_{j,t}$ and j , respectively at time t . In the literature, it is common to use either this approach or to scale the sum of total trade by the sum of both countries’ GDP. However, as argued in Frankel and Romer (1999), countries with higher income are possibly more active in trading. Therefore, using bilateral trade as a percentage of GDP as a determinant of BCS could lead to biased results as our country sample is very heterogeneous with respect to income levels.

As discussed above, the endogeneity of the relationship between trade integration and BCS has been widely acknowledged in the literature. To tackle this issue, we follow e.g. Calderón et al. (2007), Frankel and Rose (1997) or Frankel and Romer (1999) by instrumenting bilateral trade intensity with a gravity model variable. As an instrument, we use the log of the distance between the capital cities of each country pair, which is taken from the CEPII’s database on geographical variables⁷.

Another commonly used variable is the degree of economic specialization, which could affect the impact of trade on BCS. In this case, we follow Calderón et al. (2007) and Barajas et al. (2012) and calculate a simple asymmetry of production index

(IV)

$$ASP_{ij,t} = \frac{1}{k} \sum_{k=1}^3 |v_{ki,t} - v_{kj,t}|$$

⁷ The CEPII database does not provide individual data for Serbia and Montenegro. Therefore, we use the same measure for both countries.

where $ASP_{ij,t}$ denominates the mean⁸ of absolute differences of the value added share in the total production of each country v of sector k for each country pair in each year. Like Barajas et al. (2012), we focus on three sectors, namely agriculture, industry and services.

Fiscal policy is another possible determinant of BCS. On the one hand, fiscal policy can be used as a stabilizer at the national level to help smoothing the business cycle but on the other hand, fiscal policy can by itself be the source of idiosyncratic shocks. To account for the role of fiscal policy, usually the difference between a country pair's budget balances is used for measuring fiscal differences between country pairs; but as the budget balance itself is affected by the business cycle, the issue of endogeneity has to be tackled to avoid reverse causality issues. In this paper, we follow a popular approach by using the difference of the cyclically adjusted budget balances CAB between each country pair i and j in each year t .

(V)

$$FD_{ij,t} = |CAB_{i,t} - CAB_{j,t}|$$

For the EU countries, the cyclically adjusted budget balance is available from the European Commission's annual macro-economic (AMECO) database. Estimates of the cyclically adjusted budget balance do not exist for the Western Balkan countries, however.⁹ Therefore, we use the budget balance reported for each country in the Vienna Institute for International Economic Studies (wiiw) database and adjust it using the method employed by the European Commission (see Mourre et al., 2014).¹⁰

The role of common monetary policy in BCS has entered the discussion mainly with respect to the question whether the business cycles of members of currency areas tend to synchronize – either because currency areas boost trade, which in turn increases output synchronization, or because there is a currency union effect *per se* (see Gächter and Riedl, 2014). Belke and Zenkic (2007) argue that the choice of exchange rate regime plays an important role in making the Western Balkan countries' transition process, and

⁸ Instead of using the sum of absolute differences as in Barajas et al. (2012), we use the mean given that data points are missing in some sectors.

⁹ Croatia's cyclically adjusted budget balance has been available on AMECO only since 2001; in our paper, we therefore treat Croatia like the other Western Balkan economies.

¹⁰ Following the European Commission, we calculate the cyclically adjusted budget balance by $CAB_{i,t} = \frac{B_{i,t}}{Y_{i,t}} - \varepsilon * \frac{(Y_{i,t} - Y_{i,t}^p)}{Y_{i,t}^p}$ where $\frac{B_{i,t}}{Y_{i,t}}$ denotes the nominal budget balance in terms of country i 's GDP at year t . ε stands for budgetary semi-elasticity, which is a measurement of the budget balance's reaction to the level of the output gap $\frac{(Y_{i,t} - Y_{i,t}^p)}{Y_{i,t}^p}$. To calculate the output gap, we take the potential output $Y_{i,t}^p$ obtained from the HP filter described in section 2.1; for ε we assume a semi-elasticity of 0.42, which is the unweighted average of the budgetary semi-elasticities of the individual countries that joined the EU after 2004.

eventually their further EU integration, a success.¹¹ In the regression, common monetary policy is accounted for by using a bilateral dummy variable. It takes the value of 0 at all points in time for Albania, Croatia and Serbia, since these countries – at least *de jure* – do not fix their exchange rates. For the remaining Western Balkan countries, the dummy variable is set to 1 vis-à-vis Germany, starting from the year in which their currencies were pegged to the Deutsche mark (1995 for FYR Macedonia, 1998 for Bosnia and Herzegovina, 1999 for Montenegro). As of the introduction of the euro in 2002, the dummy variable takes the value of 1 vis-à-vis each country that entered the euro area in 2002, i.e. the 12 original euro area countries. Subsequently, the dummy variable for Bosnia and Herzegovina, FYR Macedonia and Montenegro changes to the value of 1 vis-à-vis Slovenia (2007), Cyprus and Malta (2008), Slovakia (2009) and Estonia (2011).

To test for the impact of financial flows, we follow Artis et al. (2008) and use FDI flows scaled by both countries' GDP, compiled from the United Nations Conference on Trade and Development (UNCTAD) FDI database and the wiiw database:

(VI)

$$FDI_{ij,t} = \left(\frac{FDI_{ij,t}^I}{Y_{i,t} + Y_{j,t}} \right)$$

Because only data on inward FDI are available for some Western Balkan countries, we use bilateral net inward FDI flows between the countries i (EU country, sender) and j (Western Balkan country, recipient), scaled by the sum of both countries' GDP. We do not expect that excluding outward FDI flows will bias our results as the available data show that outward FDI flows from the Western Balkans to the EU are negligible.

Another proxy for financial flows is taken from the Bank for International Settlement (BIS) locational¹² banking statistics database, which provides data on international financial claims and liabilities of bank offices residing in the BIS reporting countries vis-à-vis the country of residence of the bank's respective counterparty. However, reporting countries neither include the CESEE EU Member States nor Malta, which substantially constraints the observations. In a manner analogous to the construction of the trade and FDI variable, we define

(VII)

$$BF_{ij,t} = \left(\frac{A_{ij,t} + L_{ij,t}}{Y_{i,t} + Y_{j,t}} \right)$$

¹¹ The authors would like to point out that other factors such as legal and institutional reforms are relevant as well. However, these factors are largely time invariant and have already been incorporated in the regression by country-pair dummies.

¹² We use the locational instead of the consolidated banking statistics because they include lending to subsidiaries and affiliates, which would be netted out otherwise.

where bank flows between two countries i and j , $BF_{ij,t}$ are measured as the sum of exchange rate-adjusted flows in assets A_{ij} and liabilities L_{ij} divided by the sum of both countries' GDP.

Remittances are an important source of income in most Western Balkan countries. To take account of the importance of these flows to the Western Balkans, we investigate the effect remittances on BCS between the sending and the recipient country. No data are available on bilateral remittances that fully cover our country sample and time period. However, the World Bank provides data on the aggregate remittance flows the Western Balkan countries received from the EU in a specific year. We use these data for calculating proxies for bilateral flows of remittances.¹³ Bilateral migration data are provided by UNCTAD:¹⁴

(VIII)

$$R_{ij,t} = \frac{M_{ij,t}}{M_{j,t}^{EU}} * R_{i,t}^{EU}$$

Remittances R sent from country i to country j at time t are calculated by dividing the number of migrants M of home country j living in host country i at time t by the total number of migrants of the home country (migration share), which is multiplied by total remittances originating from the EU. It has to be noted that, by construction, business cycles of the respective EU countries are not reflected in the constructed variable. However, there is empirical evidence that cyclical output developments of the sending country are more or less irrelevant for the propensity to remit (Akkoyunlu and Kholodilin, 2008; Sayan and Tekin-Koru, 2012; Vargas-Silva, 2008).

For estimation¹⁵ we use a system GMM estimator developed by Blundell and Bond (1998), which allows us to instrument endogenous variables with their own lags. Additionally, we employ one external instrument (distance between capital cities) for all regressions. Asymmetry of production, fiscal differences, quasi-common monetary policy and the distance between capital cities are considered exogenous control variables or instruments, while the other variables are treated as endogenous variables because they are correlated with past and possibly current realizations of the error term. Because the two-step estimation typically yields standard errors that are downward biased, we choose the more efficient Windmeijer's finite sample correction for the two-step covariance matrix.

¹³ As inflows of remittances to the home country are significantly correlated with the degree of migrants living in the migrants' host country (see e.g. IMF, 2005, and Lueth and Ruiz-Arranz, 2008), we assume that the amount remitted to a Western Balkan country from a certain EU country depends on the number of migrants living in the EU host country as a share of the total number of migrants living in the EU.

¹⁴ Bilateral migration flows are only available for the years 1990, 2000, 2010, 2012 and 2013. For missing years, we approximated the values by calculating moving averages.

¹⁵ Our estimation was carried out in STATA, using the `xtabond2` environment developed by Roodman (2009).

3.3 Regression results

Table 2 shows the estimation results of the regressions performed. In addition to the obtained coefficients, we show the results of the Arellano-Bond tests for autocorrelation as, by construction, in first differences an autoregressive [AR](1) process is expected but autocorrelation of AR(2) should not be present because that would indicate that the second lags of endogeneous variables are poor instruments. We also include the p-values of the Hansen test for overidentifying restrictions. Column (1) reports the results from the “baseline” regression that includes bilateral trade, asymmetry of production and fiscal differences. Columns (2) to (5) add several variables, one at a time, to the baseline model, namely common monetary policy (3), bilateral FDI flows (4), bilateral bank flows (5) and bilateral remittances (6). The results of the Arellano-Bond tests confirm that the lags used as instruments are valid and the Hansen p-values indicate that the results are robust to overfitting.

Table 2

Regression results

Variable/model	(1)	(2)	(3)	(4)	(5)
Lagged bilateral business cycle synchronization	0.080*** (0.028)	0.085*** (0.028)	0.071** (0.030)	-0.013 (0.046)	0.0380 (0.030)
Bilateral trade	0.046* (0.024)	0.038* (0.021)	0.075*** (0.028)	-0.020 (0.035)	0.044** (0.020)
Asymmetry of production	0.003 (0.007)	0.005 (0.007)	-0.003 (0.009)	-0.016 0.012	0.012 (0.007)
Fiscal differences	-0.018** (0.009)	-0.018** (0.008)	-0.028*** (0.009)	0.006 (0.012)	-0.020** (0.008)
Common monetary policy		0.065 (0.071)			
Bilateral FDI			-0.009** (0.004)		
Bilateral bank flows				-0.011** (0.004)	
Bilateral remittances					-0.015* (0.008)
Constant	1.006** (0.514)	1.102*** (0.214)	0.598*** (0.177)	0.365 (0.670)	0.759*** (0.227)
Arellano-Bond test AR(1)	-8.16***	-8.20***	-7.40***	-4.33***	-7.58***
Arellano-Bond test AR(2)	1.24	1.32	0.89	-0.31	0.75

Hansen p-value	0.989	0.991	0.956	0.997	0.966
Number of instruments	205	206	192	135	192
Observations	2,245	2,245	1,994	994	1,905

Source: Authors' calculations.

Note: Dependent variable: Bilateral BCS. Standard errors are reported in parentheses. *, **, *** indicate a significance level of 10%, 5% and 1%, respectively. Out-of-sample instrument included: *logdistcap*. In-sample instruments: up to 4 lags. Time dummies are included but not reported. Maximum time span: 1994–2013.

The results of the baseline regression are in line with the findings of empirical literature. The coefficient of bilateral trade is positive as expected. As pointed out above, some authors have argued that the trade effect for uneven country pairs is usually negligible. The magnitude of the trade coefficient in our baseline regression does not support this hypothesis; however, it can be noted that the coefficient obtained is only one-third of the size of the trade coefficient identified in Gächter and Riedl (2014), who use a similar econometric model but estimate the determinants of BCS between countries within the EU. The asymmetry of the production index yields no significant result; thus we conclude that economic specialization did not play a role in the business cycle convergence process between the Western Balkans and the EU in the observation period. In the baseline regression we also test the effect of fiscal differences; the resulting negative coefficient is in line with the majority of earlier studies (see e.g. Darvas et al., 2005; or Crespo-Cuaresma et al., 2011) which argued that fiscal policy was the source of idiosyncratic shocks and thus led to greater business cycle divergence.

In the next step (2), we add a dummy variable for common monetary policy as described above. However, using the dummy variable does not yield any statistically significant result, meaning that *ceteris paribus* a Western Balkan country that uses the euro as its currency or nominal anchor does not exhibit higher business cycle correlation with the euro area countries than other Western Balkan countries.

In column (3) we add FDI to test for the effect of financial flows on BCS. The result confirms the negative impact of financial flows found by Kalemli-Ozcan et al. (2009) and García-Herreo and Ruiz (2008). Our result points to the argument that FDI flows are procyclical and thus reduce BCS between the Western Balkans and the EU. Although the coefficient seems small, it is rather large when we look at standardized coefficients, which is useful when the regressors are scaled differently (Wooldridge, 2009), as is the case for trade and FDI in our sample. The standardized coefficient¹⁶ is 0.130 for trade and -0.182 for FDI; thus, an increase in the standard deviation of FDI has a stronger effect on BCS than an increase in the standard deviation of trade. A comparison of the coefficient of trade in model (3) with the coefficient obtained in the baseline model (1) supports the

¹⁶ The standardized coefficient is calculated by multiplying the derived coefficient by the standard deviation of the respective independent variable and dividing it by the standard deviation of the dependent variable (i.e. BCS). The transformation of trade is thus obtained by $0.075 * \frac{2.028}{1.174} \approx 0.130$; the transformation of FDI by $-0.009 * \frac{23.689}{1.174} \approx -0.182$.

argument brought forward by Déés and Zorell (2011) that FDI has an indirect positive effect on BCS via trade.

In model (4) we attempt to test whether the negative correlation coefficient of FDI obtained in model (3) holds when another measure of financial flows is used, namely the flows of bank assets and liabilities between two countries. Unfortunately, because data are not available for all CESEE EU countries and for Malta, about half of the observations have to be dropped in the regression, which causes almost all indicators to become insignificant. However, the coefficient of bank flows is significant, negative and about the same size as the coefficient of FDI in model (3), which tends to confirm the negative impact of financial flows on BCS.

Column (5) reports the regression results obtained when taking into account bilateral remittance flows. In contrast to the results of Barajas et al. (2012), the obtained coefficient is negative, indicating that remittance flows from the EU to the Western Balkans decrease BCS. To the best of our knowledge, no theoretical or empirical research apart from Barajas et al. (2012) has so far included the role of remittances in BCS. However, there is literature that investigates the relationship between remittances and the business cycle of the recipient country.¹⁷ There are two hypotheses on the motives behind remittances (Chami et al., 2008). On the one hand, when sent in order to take advantage of high returns or favorable economic conditions, remittances could exhibit procyclical properties (see Sayan and Tekin-Koru, 2012, for Turkey; Isokovic and Ilgun, 2015, for Bosnia and Herzegovina; Lueth and Ruiz-Arranz, 2006, for a global data set). If this was the case, remittance flows would have a negative effect on BCS by enhancing the size of the business cycle of the recipient country, similar to financial flows. On the other hand, remittances could be used to compensate recipients for unfavorable economic conditions and thus help smooth their consumption patterns. This hypothesis would imply anticyclical behavior with respect to the business cycle (see e.g. Frankel, 2011; Sayan, 2006) and a positive impact of remittances on BCS between the sending and the receiving country.¹⁸ Our result suggests that remittances from the EU to the Western Balkans exhibit procyclical behavior with respect to the business cycle in the Western Balkan economies. Standardizing the coefficient of remittances yields a value of 0.036, while the standardized coefficient of trade in this regression is 0.076.¹⁹ Thus, the effect of an *increase* in a standard deviation of remittances on BCS is about half the size of the effect of a *decrease* in a standard deviation of trade.

¹⁷ As discussed above, empirical evidence points to the fact that cyclical output developments of the sending country are more or less irrelevant for the propensity to remit.

¹⁸ Motives behind remittances do not necessarily have to be congruent with their final use; microeconomic evidence suggests that the majority of remittances is often used for consumption, but that a significant part of remittances also goes into saving or investment (Chami et al., 2008).

¹⁹ The transformation of trade is obtained by $0.044 * \frac{2.028}{1.174} \approx 0.076$; the transformation of remittances by $-0.015 * \frac{2.790}{1.174} \approx 0.036$.

According to the literature, it may be possible that idiosyncratic shocks are transmitted with lags; if this was the case, explanatory variables in one period would affect BCS mainly in the subsequent period. To test this assumption and to put additional restrictions on reverse causality issues at the same time, the most important regressions provided in table 3 are re-estimated by lagging all explanatory variables by one period (the lagged bilateral dependent variable is kept at the lag of one period).

Table 3

Regression results when lagging explanatory variables

Variable/model	(1)	(3)	(5)
Lagged bilateral business cycle synchronization	0.070** (0.028)	0.046 (0.028)	0.041 (0.030)
Lagged bilateral trade	0.049** (0.020)	0.076*** (0.024)	0.059*** (0.022)
Lagged asymmetry of production	-0.001 (0.007)	-0.013 (0.009)	0.003 (0.007)
Lagged fiscal differences	0.042*** (0.013)	0.046*** (0.015)	0.048*** (0.011)
Lagged bilateral FDI		-0.009*** (0.003)	
Lagged bilateral remittances			-0.024* (0.013)
Constant	0.989 (0.762)	0.367** (0.182)	1.647** (0.651)
Arellano-Bond test AR(1)	-8.11***	-7.33***	-7.35***
Arellano-Bond test AR(2)	0.82	0.05	0.20
Hansen p-value	0.957	0.934	0.944
Number of instruments	196	182	182
Observations	2,127	1,878	1,760

Source: Authors' calculations.

Note: Dependent variable: Bilateral BCS. Standard errors are reported in parentheses. *, **, *** indicate a significance level of 10%, 5% and 1%, respectively. Out-of-sample instrument included: *logdistcap*. In-sample instruments: up to 4 lags. Time dummies are included but not reported. Maximum time span: 1995–2013.

The robustness test of lagging the explanatory variables by one year yields rather interesting results. The signs of coefficients of trade, FDI and remittances do not change; instead, the magnitude of the coefficient of trade and of the coefficient of remittances

even grows. This leads to the conclusion that idiosyncratic shocks are indeed transmitted with lags. In contrast to the other explanatory variables, the coefficient of fiscal differences becomes positive, larger in magnitude and more significant. The regression results presented in table 3 show that fiscal differences are negatively correlated with BCS in the same year. Thus, we argued that fiscal differences are the result of idiosyncratic shocks. However, the outcome of lagging the explanatory variables suggests that fiscal differences actually lead to *more* BCS in the following year. The positive sign of the coefficient even holds when the fiscal difference variable is lagged by two years.²⁰ As argued above, while we cannot instrument fiscal differences, we use the cyclically adjusted budget balance to correct the cyclical component of fiscal spending. However, regressing the fiscal differences on the BCS of the same year does not completely solve the issue of reverse causality. If policymakers anticipated a recession (boom) during the current year or even before the year begins, they could increase (decrease) public consumption or investment to smooth the business cycle. Under the assumption that anticyclical policy needs some time to become effective, such a policy reaction would suggest a negative relationship between BCS fiscal differences in the same year, but as soon as the anticyclical measures unfold, fiscal differences would lead to higher synchronization in subsequent years. To conclude, the changing sign of fiscal differences suggests that fiscal policy is used as an economic stabilizer to help smoothing the business cycle rather than being the source of idiosyncratic shocks.

Additionally, to rule out that the results obtained are driven by the dynamics of one country we run the regressions of table 2 again, excluding one Western Balkan country at a time. The estimated coefficients prove to be robust in the sense that the signs do not change.²¹ However, some of the coefficients become insignificant, which can be traced back to the loss of observations (25 country pairs multiplied by the available time span) in each regression. Similarly, we also estimate each of the models in table 2 with a smaller data set that comprises only the current 19 euro area countries instead of the EU-25. Again, our results do not change qualitatively.

4 Conclusions

This paper fills several gaps in the literature. It is the first that investigates the business cycle synchronization (BCS) of *all* Western Balkan economies with the EU-25, i.e. the EU excluding Denmark, Croatia and the U.K., and the first that empirically identifies the determinants of BCS between the two regions. For this purpose, we use a period-by-period correlation index to analyze the convergence process on a yearly basis. Because BCS estimations are prone to endogeneity problems, we employ a system GMM estimator that instruments potentially endogenous variables with their own lagged values.

We clearly identify a process of business cycle convergence between the Western Balkan economies and the EU-25 aggregate from the early transition phase in the 1990s up to the

²⁰ For brevity reasons, results are not reported here but are available from the authors upon request.

²¹ For brevity reasons, results are not reported here but are available from the authors upon request.

year 2013. While prior to 2009, convergence was higher for Bosnia and Herzegovina, Croatia, FYR Macedonia and Montenegro than for the other Western Balkan countries, after the 2009 crisis year BCS has been high for all Western Balkan countries. Thus, in recent years, the lack or narrow limits of independent monetary policy in the Western Balkans seem not to have been very costly from the perspective of business cycle developments. However, it remains to be seen whether the high degree of business cycle convergence will continue in the longer term. Moreover, it should be recalled that in this study we only examine one of the multiple OCA criteria.

With respect of the determinants of BCS between the Western Balkans and the EU, we find that foreign trade is the most important positive factor. This result is in line with earlier studies that used different regional or global samples. Another variable that is usually tested for in the BCS literature are the fiscal differences between two countries. Here, our results differ somewhat from the findings of other papers. While we also find a negative coefficient of fiscal differences for BCS in the same year, the sign of the coefficient becomes positive in the subsequent years. It therefore seems that fiscal policy is used as an economic stabilizer to help smoothing the business cycle rather than being a source of idiosyncratic shocks as has been argued in earlier studies.

We also include two possible determinants of BCS that are less frequently used in the literature but which we assume to be very important in our specific country sample. One is FDI inflows, which poured into the Western Balkans especially in the years preceding the crisis. While empirical studies have not delivered a definite answer on the effect of financial flows on BCS so far, our results show that in the case of the Western Balkans, FDI has led to business cycle divergence. We explain this outcome by the procyclical nature of FDI.

There is hardly any literature on the impact of remittance flows on business cycle synchronization. One study discovered a positive impact of remittances on BCS. Our results are to the contrary, as we find that remittances sent from EU countries to the Western Balkans actually lead to business cycle decoupling. This supports the hypothesis that remittances exhibit procyclical properties with respect to the receiving economy, similarly to FDI flows.

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Annex

Table A1

Data sources

Variable	Source
Real GDP	WEO
Trade	IMF DOTS Database
Gravity variable (distcap)	CEPII
Asymmetry of production	WDI
Cyclically adjusted budget balance	Ameco, wiiw
FDI	UNCTAD FDI, wiiw
Bank flows (BF)	BIS statistics
Remittances	World Bank
Migration data	UNCTAT

Source: Authors' compilation.

Table A2

Descriptive statistics

Variable	Number of observations	Mean	Standard deviation	Minimum	Maximum
BCS	2,801	1.037	1.174	-1.289	7.991
Log(trade)	2,402	-8.101	2.028	-16.888	-2.694
Asymmetry of production	2,986	9.122	6.157	0.047	38.139
Fiscal differences	2,554	3.579	3.473	0	51.149
Monetary policy	3,750	0.141	0.348	0	1
Log(FDI)	2,146	-34.846	23.689	-57.565	-4.501
Log(bank flows)	1,244	-35.229	23.829	-57.565	-3.766
Remittances	2,275	0.842	3.007	0	27.745

Source: Authors' calculations.