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LABOUR MIGRATION IN THE EUROPEAN UNION: THE CASE OF CENTRAL AND EASTERN EUROPE

ABSTRACT: *This paper examines migration trends in the European Union since the enlargements of 2004–2007, which brought 100 million citizens of 11 Central and Eastern European countries into the EU. We examine country- and regional-level data on migration trends and show how European integration depleted the labour force in the new member countries. Several of them have lost 10% of their population since 2006, most of it via negative net migration. In 2019, 18% of Romanians, 14% of Lithuanians, 13% of Croats, and 13% of Bulgarians lived in another EU country. The quantitative analysis shows that migration contributed positively to regional convergence, as every percentage point of net migration increased GDP per capita by roughly 0.01% and reduced unemployment by 0.1–0.2 percentage points. To disentangle aggregate migration effects, further analysis will be needed to quantify its impact on regions that lose their population via migration.*

KEY WORDS: *migration, labour markets, convergence, European Union.*

JEL CLASSIFICATION: F22, F66, J61, O15, R11, R23

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1. INTRODUCTION

With borders closed due to the Covid-19 crisis, and movement complicated by medical tests and concerns about incompatible healthcare systems, migration collapsed in 2020, across Europe and globally. Yet, as recently as January 2020, the political scientist Ivan Krastev had declared (out)migration to be Eastern Europe's biggest problem.[1] Migration has been blamed for the rising popularity of fringe political parties in Eastern Europe and also in France, Sweden, and Italy. Indeed, the decision to leave the EU by Great Britain was motivated by the desire to 'take back control', mainly over migration.[2]

The real impact of migration is multifaceted. It has economic effects on growth, wages, and unemployment; social aspects such as overwhelmed healthcare facilities and schools with a rising share of children not speaking the official language; and, perhaps most importantly, cultural effects when people fear being threatened by unknown languages and immigrants' traditions.

This paper focuses on the economic aspects of migration, namely its effect on economic growth and labour markets. We use the EU enlargement of 2004–2007, which expanded the EU labour market by including more than 100 million citizens from 11 Central and Eastern European (CEE) countries: Bulgaria, Croatia, Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Romania, Slovakia, and Slovenia. We show that the migration effects in these countries were often overwhelming. Lithuania and Latvia have lost 10% of their respective populations since 2004, while 18% of Romanians now live outside their country.

We further examine how the massive reallocation of the labour force affected the convergence of living standards across the EU. There are several channels through which migration works: a more effective allocation of labour to high-productivity regions should raise the total EU growth rate and limit wage increases, while its impact on unemployment depends on the labour market structure and skill mismatch between local labour markets and migrants. However, a careful analysis of these effects requires more granular data than

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1 Ivan Krastev: Depopulation is Eastern Europe's Biggest Problem. Financial Times, 27 January 2020.

2 We include the UK in our study, as it was an EU member during the study period, and a major receiving country of net migration.

country-level migration numbers. To that end, we gather and use data from 268 NUTS2 regions – regions defined by the European Union for statistical purposes. While these regions are heterogeneous in size, population density, income, and skills, they allow quantitative analysis of migration effects on a much more detailed level than country-level data.

Using models for real income convergence and unemployment rates, we show that net migration contributes to income convergence by as much as 0.1% per percentage point in net migration. These estimates are higher than in most of the literature from the 2000s, mostly owing to the larger sample. Similarly, our estimates affirm that unemployment is strongly path-dependent. Most importantly, the impact of net migration on the unemployment rate is found to be consistently negative, while previous studies were unable to find a statistically significant effect.

The rest of the paper is organized as follows. Section 2 discusses the literature on the effects of labour migration on economic convergence, paying particular attention to the European Union. Section 3 analyses models of economic convergence that use migration as an explanatory variable. Section 4 presents our dataset, while section 5 illustrates the main migration trends at the country level. Section 6 demonstrates NUTS2 regional data and presents its main characteristics. We discuss the quantitative models of migration and main econometric results in Section 7 and conclude with some general observations in Section 8.

2. ECONOMIC CONVERGENCE AND LABOUR MOBILITY

An extensive literature exists on the effects of migration among regions either within a country (US) or a single market (EU). Most of the literature attempts to estimate the extent to which migration alleviates regional disparities. Barro and Sala-i-Martin (1992) provide a simple framework with homogeneous labour in a neoclassical growth model. Migration to high-income regions lowers capital intensity in the rich regions, and as the labour has the same characteristics and there are no barriers to factor mobility in their model, the labour moves from low-income to high-income regions and accelerates income convergence. The receiving region's capital-to-labour ratio initially decreases, reducing productivity. However, as the Harberger model shows, lower wages will lead to higher returns to capital and will attract more investment, which will restore the capital-to-labour ratio and productivity. As both labour and capital stocks increase, the receiving regions achieve a higher steady state.

However, once the assumption of homogeneous labour is relaxed, the effects of migration are more ambiguous. Etzo (2008) suggests that heterogeneous labour may offset the scale effect of migration through the change in the ratio of skilled to unskilled workers. As a result of increased migration, disparities in income per capita at the regional level may increase, although migration allows workers to maximize their individual utility (Fratesi and Riggi, 2007). Docquier and Rapoport (2012) show that if migrants possess higher human capital and skills than stayers in the sending regions, their exit lowers the steady state in the sending regions: lower available human capital requires a lower investment rate in these regions, and the adverse effects of lower investment may outweigh the positive effects of outmigration on wages. The emigration may then slow down wage growth, and the overall growth rate decrease in the sending regions. Indeed, Docquier and Rapoport show that in the decade up to 2000 more than 40% of migrants from the sub-Saharan region and 45% of migrants from all low-income countries were highly skilled, significantly diminishing potential growth rates in the countries they left.

Kaczmarczyk (2010) illustrates the same phenomena in the migration of high-skilled Poles to Great Britain after the 2004 EU enlargement, which extended the freedom of movement to Central and Eastern European countries. While before the enlargement the migrants were predominantly low-skilled, the share with tertiary education increased by a third, to 20% of a much higher number of migrants. Ostbye and Westerlund (2006) identify a similar 'brain drain' effect when estimating the growth effects of migration in Norway and Sweden. Migration from the sending country (in this case Norway) seems to dampen convergence, while its effects in Sweden are inconclusive.

Kaczmarczyk (2010) even argues that rather than a 'brain drain', emigration from CEE countries achieved a 'brain waste', whereby highly skilled workers (admittedly measured by graduating from a university with no data on the actual quality of the education) ended up in low-skilled jobs in Western Europe, primarily Great Britain.[3] This may decrease GDP growth per capita both in the sending country (by lowering its average human capital) and in the receiving country. Freideberg (2001) studies the effect of emigration from the Soviet Union to Israel and finds likewise that emigrants are concentrated in low-skilled jobs that do not correspond to their education level.

However, an endogenous change in the technology used by an industry employing recent migrants may result in changes in productivity. Dustmann (2008) argues that an increase in the supply of unskilled workers can stimulate labour-intensive production methods (for example, agriculture specializing in more labour-intensive crops). He estimates that about two-thirds of labour market adjustments are affected by technological change.

As this brief discussion suggests, the ultimate effect of migration on unemployment, growth, and productivity depends on a number of factors, including human capital distribution, the elasticity of labour supply in receiving and in sending countries, the elasticity of substitution between native and migrant workers, national wage-determination institutions, and possibly many more (Huber, 2012; Borjas, 2003). With no firm analytical conclusions, we need to turn to empirical studies to determine the likely effects of increased migration.

Empirical studies

The empirical literature on the effects of migration was initially concerned with internal migration in the US, as it represents a large labour market with significant migration flows. Longhi et al. (2006) conclude that, on average, a 1% increase in immigration to a state within the US reduces native employment by only 0.02%. In an extensive meta-analysis of the literature, Ozgen, Nijkamp, and Poot (2010) conclude that an increase in the net migration rate of one percentage point increases the GDP per capita growth rate by 0.13% on average, but that the effects of migration remain an ongoing research issue. They summarize their

3 However, the data on the education profile of emigrating Poles and other Central and Eastern Europeans may be misleading, as the determining factor in the decision to leave seems to be youth rather than qualifications. Younger cohorts tend to have a higher share of graduates from tertiary education institutions.

meta-study around the most common econometric specification of the migration effects on economic growth and convergence:

$$\ln(y_{i,t}) = \alpha + (1-\beta)\ln(y_{i,t-1}) + \gamma(migration_{i,t}) + \delta \ln(X_{i,t}) + \varepsilon_{i,t} \quad (1)$$

The dependent variable $y_{i,t}$ is the annual growth rate per capita in a region *i* in year *t*. In this specification, β is the annual rate of beta convergence at which a region converges to its own long-run steady state, and γ is the annual net migration rate coefficient. The coefficient of the net migration variable γ estimates the impact net migration makes on the convergence. Mankiw et al. (1992) use investment rate and education profile as in the Solow model, while in their analysis of regional convergence in the EU, Fidrmuc et al. (2019) add the natural population growth rate plus the sum of technological progress and depreciation.[4]

According to Sala-i-Martin et al.'s (2004) paper, in which they estimate growth regressions separately for the US, Japan, and 5 European countries during 1950–1990, the lower capital intensity results in a lower growth rate in the destination regions and faster growth in the sending regions. However, according to their results, migration plays only a marginal role in the convergence process. Similarly, weak or insignificant effects of migration are found by Cardenas and Ponton (1995) for Colombia and Gezici and Hewings (2004) for Turkey in the 1990s. Inconclusive estimators may be a consequence of the non-linear impact of migration. Ozgen (2010) lists two major impacts of labour migration: the scale (size) effect and the composition effect. Along with Docquier and Rapoport (2012), he argues that an intensive outward migration of skilled labour diminishes productivity in the sending regions while benefiting the receiving regions with an upward shift in productivity.

Mattoo (2008) and Huber (2012) analyse the effects of migration in the receiving country. They argue that a pool of labour increased by migration should positively affect productivity and that the different skills that migrant labour possesses enhance technology adoption. Mas et al. (2008) show a small positive

4 Most studies follow Mankiw (1992) and use 0.05 as the sum of technological progress and depreciation. Fidrmuc (2019) uses 0.06 for technical reasons, to offset the negative population growth of 5% in several regions during the 2000s.

impact of labour (in)flows on the growth rate in the UK and a significantly negative impact of the outflow of workers from Spain during the 1990s and early 2000s.

The EU Experience

The European experience became more relevant after the Maastricht Treaty of 1992 established freedom of movement for EU citizens and as the poorer countries in Central and Eastern Europe began to integrate with the EU (Haas et al., 2019). East-west migration became the primary concern after the 2004–2007 EU enlargement brought ten Central and Eastern European countries into the EU, followed by an eleventh in 2013.[5] The first papers focused on the UK experience. The UK was one of the three EU countries that liberalised their labour markets for CEE immigrants in 2004 and witnessed a significant increase in inflows soon afterward.

Dustman et al. (2005) and Lemos and Portes (2008) find a small negative effect on the employment of the semi-skilled, young, and elderly in Great Britain, as immigrants typically compete with incumbents only in marginal labour market segments. Dustman is among the first to use a fixed-effects instrumental variables estimation to account for the endogeneity of migration. Blanchflower et al. (2007) estimate that more than 0.5 million migrants from the CEE region moved to Great Britain in the two years after the first wave of enlargement in 2004. He claims that the generally lower wage demands of recent migrants helped contain inflation in the UK. At the same time, the UK government expected that more than half of these migrants would return to their home countries, which mostly did not happen.

Apsite, Krisjane, and Berzins (2012) illustrate the scope of emigration in the example of Latvia: emigration from Latvia to the UK more than doubled in 2009, while in Latvia real GDP declined by a staggering 14%. The structure of departing workers changed as well: after the financial crisis of 2009 the emigrants were more likely to be younger, more-educated, and more urban than before. More than 40% of emigrants from Latvia had tertiary education; however, only 20% were able to

5 Czechia, Estonia, Hungary, Latvia, Lithuania, Poland, Slovakia, and Slovenia joined the EU in 2004 (along with Cyprus and Malta), followed by Bulgaria and Romania in 2007; Croatia then joined in 2013. These 11 countries constitute the CEE group refered to in the text.

find a job in non-manual or low administrative positions. This conclusion is supported by Simionescu (2016) and Mihi-Ramirez (2013), who show that migration flows are highly sensitive to domestic growth rates in both the CEE regions and Spain.

Barrel et al. (2010) demonstrate that net migration depends on GDP per capita (inversely) and unemployment rates. They use data from the early years of the enlargement and limit their contribution to a qualitative analysis. Anacka and Okolski (2010) show that age is an essential determinant of migration, as younger people tend to be more mobile. Marques (2010) approaches migration much more vigorously, estimating a gravity model with 15 exogenous variables. She shows that EU membership leads to an increase in migration from new member states to 'old' Europe.

While there is substantial country evidence of the effects of migration on unemployment and wages, literature with a broader European perspective remains limited. Brücker (2009) argues that migration from the new member states yields substantial gains for the GDP of an enlarged EU in the long run and that migrants themselves are the main beneficiaries of free movement. He further states that the effects on natives in sending and receiving countries are ambiguous and, in general, relatively small. Huber and Tondl's (2012) study on convergence in the EU is the most comprehensive; they estimate three models: income growth, unemployment, and productivity. While they find no significant effect of migration on unemployment rate convergence, Huber and Tondl identify a significant and positive impact of migration on GDP per capita and on productivity growth in both receiving and sending regions. They estimate that a 1% increase in immigration increases GDP per capita by 0.02% and labour productivity by about 0.03% in the immigration regions, with similar estimates for the emigration regions. The authors argue that migration can be viewed as a transfer of human capital to immigration regions that happen to have above-average GDP per capita. In this sense, migration contributes to regional divergence.

In a later paper, Huber (2018) analyses migration flows during and after the 2008–2009 recession and shows that the labour markets helped mitigate the impact of the recession, but that heterogeneity across countries and demographic groups has increased. According to Huber, after the crisis (2011–2014) the impact of migration on regional labour market adjustment appears to decline significantly.[6] Thus, the pre-crisis period seems to have been an exceptional period in terms of labour market adjustment in the EU when the labour markets of the new member states were integrated in the EU-wide migration flows.

King and Okolski (2018) look at recent migration flows in Europe in a long-term context and argue that the current intra-EU migration among member states amounts to only 0.3% of the entire population, a fraction of the US inter-state migration of 2.4%. However, they show that (e)migration from the six most active countries – Bulgaria, Croatia, Latvia, Lithuania, Poland, and Romania – has been consistently high, reaching a cumulative 9% for Romania. As our discussion in the following chapters shows, the proportion of emigrants increased further until 2019, reaching double digits in 5 CEE countries: Bulgaria, Croatia, Latvia, Lithuania, and Romania.

Kahanec and Zimmermann (2009) focus on the effects of the 2004 and 2007 enlargements of the EU and show that while increased migration flows had a significant impact on migration flows from new to old member states, any adverse effects on wages or employment in the labour market in the receiving countries were insignificant. Economic migration should result in a more efficient allocation of production factors, thus improving economic growth prospects. Migration also contributes to the transfer of knowledge and technology, which may lead to a one-way ‘brain drain’ from less developed to more prosperous countries. In a more benign scenario, migration may result in two-sided ‘brain circulation’ between the host country and the country of emigration. However, the authors are among the first to highlight rising labour market pressures in sending regions, quoting labour shortages in Lithuania and Poland.

The literature seems to converge in the view that migration is favourable for both the sending and receiving region when the brain drain is limited and migration is not one-way but mutual. Whether or not the total outcome is positive remains an issue that calls for careful estimation.

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6 Our analysis seems to confirm this view – see Chapter 4.

3. GROWTH MODELS WITH MIGRATION

Labour flows within the European Union have become one of the main instruments of improving the individual well-being of migrants from new member states. To analyse the impact of migration on the convergence in real GDP, unemployment rates, and productivity we follow a standard setup suggested by many studies, including Borjas (1999), Ozgen (2010), Huber (2012), and Wolszczak-Derlacz (2009). We test whether migration significantly impacted changes in these three variables at the level of NUT2 regions in the European Union. Our data cover the 2006–2018 period, for which net migration rates are available.

$$\begin{aligned} &\ln\left[\frac{GDPPC_{i,t}}{GDPPC_{i,t-1}}\right] = \alpha + \beta\ln(GDPPC_{i,t-1}) + \gamma(migration_{i,t}) + \delta\ln(X_{i,t}) + \varepsilon_{i,t} \\ &\ln\left[\frac{U_{i,t}}{U_{i,t-1}}\right] = \alpha + \beta\ln(U_{i,t-1}) + \gamma(migration_{i,t}) + \delta\ln(X_{i,t}) + \varepsilon_{i,t} \\ &\ln\left[\frac{PROD_{i,t}}{PROD_{i,t-1}}\right] = \alpha + \beta\ln(PROD_{i,t-1}) + \gamma(migration_{i,t}) + \delta\ln(X_{i,t}) + \varepsilon_{i,t} \end{aligned} \tag{2}$$

As dependent variables, $GDPPC_{i,t}$ is the growth rate of per capita GDP in purchasing power parity, $U_{i,t}$ is the unemployment rate, and $PROD_{i,t}$ is productivity approximated by GDP per active workers. The $migration_{i,t}$ variable measures net migration to/from the region. We employ a set of control variables $X_{i,t}$ that typically includes demographic and educational variables, labour market characteristics, and the investment rate in the case of income (*GDPPC*) and productivity (*PROD*) equations. All variables are annual and structured by NUTS2 regions as balanced panel data (more on the data in the next section).

The equations can be transformed to a linearized form (3) that allows an estimate of the convergence speed:

$$\begin{aligned} &\ln\left(GDPPC_{i,t}\right) = \alpha + (1-\beta)\ln(GDPPC_{i,t-1}) + \gamma(migration_{i,t}) + \delta\ln(X_{i,t}) + \varepsilon_{i,t} \\ &\ln\left(U_{i,t}\right) = \alpha + (1-\beta)\ln(U_{i,t-1}) + \gamma(migration_{i,t}) + \delta\ln(X_{i,t}) + \varepsilon_{i,t} \\ &\ln\left(PROD_{i,t}\right) = \alpha + (1-\beta)\ln(PROD_{i,t-1}) + \gamma(migration_{i,t}) + \delta\ln(X_{i,t}) + \varepsilon_{i,t} \end{aligned} \tag{3}$$

The specifications (3) may suffer from an endogeneity problem, as migrants are attracted to regions with higher income (GDP) and/or with lower unemployment (Borjas, 2001). In the first approximation the panel data analysis with fixed and random effects deals with this problem (Ozgen, 2010). This may lead to reverse causality and unrealistic high estimators for migration coefficients in a panel data regression with fixed effects across the regions, the so-called Nickell bias (Nickell, 1981). The endogeneity problem is typically mitigated by implementing instrumental variables in the form of lagged migration rates, using dynamic panel estimates with fixed/random effects. The Hausmann test excluded the random effects method as potentially inconsistent and biased, so we report only the fixed effects results (Hausman, 1978).[7]

Additionally, we were unable to distinguish between international and internal migration, as was done by Huber (2012). In his paper, Huber shows that international migration tends to give more consistent and higher estimators, while domestic migration within a country typically has little effect on dependent variables. Eurostat, however, does not separate migration data into international and internal migration, which may lead to lower estimators in our specification.

4. DATA

Our dataset for 2005–2018 covers all 28 countries that were members in 2019 (i.e., including the UK and all countries that joined in 2004–2013). We use annual data structured by the European Union's NUTS2-level regions, which provides a much more granular and rich view of the convergence process in the EU than a national-level analysis. The NUTS2 standard was formally adopted in 2003 and was revised in 2006, 2010, 2013, and 2016, representing a challenge for creating consistent data series. There were 281 NUTS2 regions in the EU, including the UK, between 2006 and 2018, the final year of our sample, with several adjustments, most prominently in France.

Most regional data is available from Eurostat (Regional Statistics by NUTS Classification), with additional data from the Annual Regional Database of the

[7] To check for remaining endogeneity, Blundell and Bond (1998) recommend generalised method of moments (GMM) with lagged levels and differences as instruments. In our setting, the GMM method proved to be unstable due to the extremely high number of instrumental variables.

European Commission's Directorate General for Regional and Urban Policy (ARDECO). Eurostat data include GDP data in nominal, real, per capita, and purchasing power parity terms. Gross fixed capital formation, employment, and compensation are available from Eurostat. Additional labour market variables available at ARDECO are also used. Demographic data, including crude rates of total population change, natural change, and net migration, are published by Eurostat. Labour market data is available in the regional detail at Eurostat, and we use its data on total employment, unemployment rate, the share of long-term unemployed, and hourly wages. To approximate skill level at the regional level we use data on the share of workers with tertiary and secondary education, available at Eurostat.

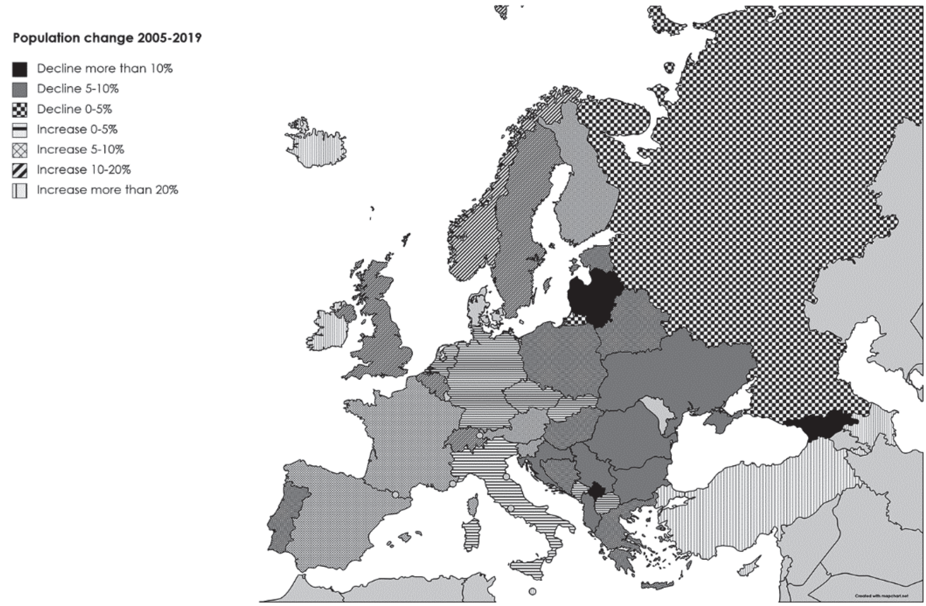
Data availability has improved significantly since papers published in 2009–2012, but several gaps persist. First, we excluded 5 French extraterritorial departments and 2 Spanish regions located in North Africa, as all these are too small and too distinct from the remaining regions. The data was also insufficient to include the smallest Finnish region, Swedish-speaking Aland Island (roughly 30,000 inhabitants). Two NUTS2 regions in former Eastern Germany, Leipzig and Chemnitz, did not provide migration rates before 2011 so we had to exclude them from the sample as well. In the UK, London was split into five NUTS2 regions, which did not provide migration and unemployment data until 2012. However, we were able to use data from the two original NUTS2 regions instead. In Slovenia and Croatia the unemployment rate was not reported by the countries' two NUTS2 regions, so we used the national level for Croatia until 2008 and for Slovenia until 2010. Denmark only defined its NUTS2 regions in 2007 and did not report data on the share of tertiary and secondary education by NUTS2 regions. Instead, we used national data for the country's educational stratification, and we approximated migration data for 2006.

Our data set is more than double the size of the most extensive analysis of regional convergence, Huber (2012), as we use 3,752 observations for the 2005–2018 period. The migration data are available for 2006–2018 period only, limiting our sample to 3,484 observations. Due to insufficient data, Hubner excluded all regions from Bulgaria and the UK, i.e., the two most active sending and receiving migrant countries, respectively.

5. COUNTRY-LEVEL STYLIZED FACTS

The EU population increased to 513.5 million in 2019, from 501 million at the outset of the great recession in 2009 and 490 million at the eve of enlargement in 2004. However, while the population has increased in 18 countries since 2004, it has declined in 10 countries. Only two of the latter ten are 'old' member states: traditionally labour-exporting Portugal and Greece. [8] Eight countries with declining populations are from the CEE region: Bulgaria, Croatia, Estonia, Hungary, Latvia, Lithuania, Poland, and Romania. Only 3 of the 11 CEE countries – Czechia, Slovakia, and Slovenia – increased their population, by a modest 2%–3%. Latvia and Lithuania (and non-EU Ukraine) were most affected in relative terms, losing almost a fifth of their respective populations since 2004, with Bulgaria and Romania losing 10% (see Figure 1).

Figure 1: Population change in Europe, 2005–2019

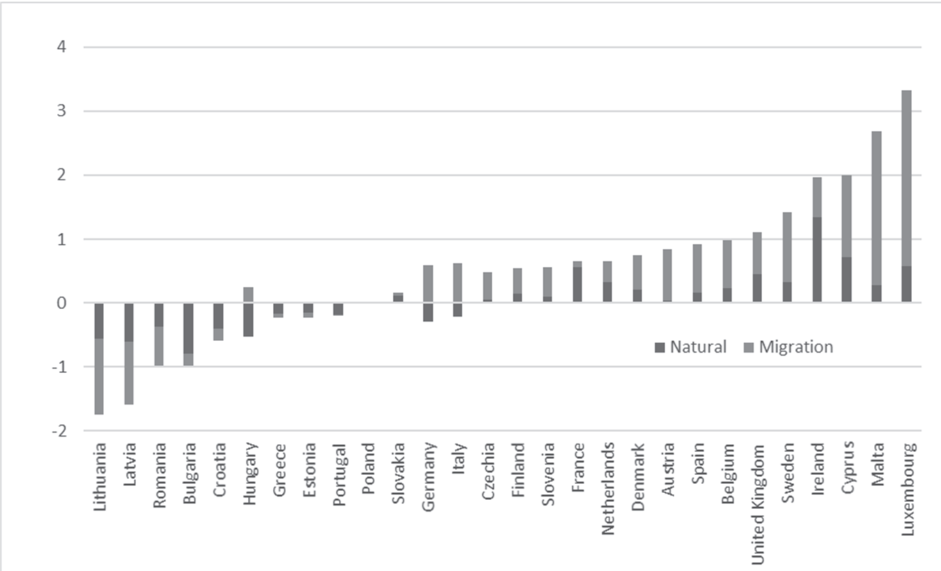


Source: Eurostat database, Annual Regional Database, author's calculation.

[8] For the purpose of our discussion,we include Cyprus and Malta, which joined the EU among 'old' members in 2004. Their combined population of 1.2 million represents 0.2% of the EU population and thus does not impact our conclusions.

Migration flows within the European Union became the defining characteristic of the post-crisis 2010s.[9] In the CEE countries, migration has been blamed for labour shortages and the rise of extremism. It is also held responsible for intensifying the anti-EU feelings in the UK that contributed to Brexit (Blinder, Richards, 2020). De-population in the CEE region (and generally in Eastern Europe) is driven by both natural population change and net migration. However, the migration effect was particularly strong in the Baltic region and the Balkans (Figure 2). Positive net migration in Czechia, Hungary, Slovakia, and Slovenia came mainly from other CEE countries and/or from Ukraine, which has lost more than 5.5 million inhabitants since 2004, with most moving to Poland, Slovakia, Hungary, and Czechia. In the period 2004–2019 the natural population change was mildly positive only in Czechia, Poland, Slovakia, and Slovenia.

Figure 2: Natural population change and net migration in Europe, 2005–2019



Source: Eurostat database, Annual Regional Database, author's calculation.

Note: In per cent

[9] Migration from non-EU countries such as Ukraine and Moldova is even larger in relative terms (Figure 1), but regional data for these countries are much less accurate so we focus on the EU member countries only.

The three enlargement waves increased the EU's population by 25%; i.e., 100 million people. The EU labour force increased by a fifth, to roughly 250 million. However, only 3 out of 15 original EU member countries – the UK, Ireland, and Sweden – opened their labour markets from the outset of enlargement, with other countries imposing restrictions on their labour markets for a maximum of seven years. The European Commission had initially reported that the enlargement would have a relatively limited impact on labour markets within the EU (European Commission, 2008). Inflows of workers from the EU-8 countries, as it was then, increased from around 1 million in 2004 (0.2% of the total EU-28 population) to 2.3 million in 2010 (0.5% of the EU-28 population).

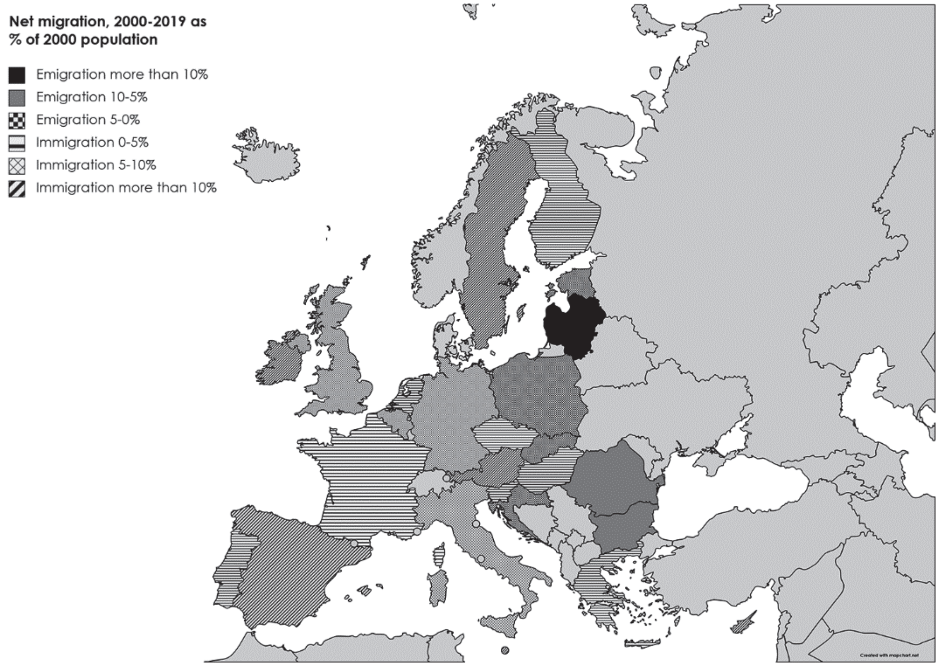
Further enlargements in 2007 and 2013 changed the migration landscape significantly. In 2018 there were more than 9 million citizens from the EU-11 living in the old EU-15 plus Cyprus and Malta. More than half of total EU citizens living in another EU country were from CEE. Romania had the largest diaspora: more than 3.5 million Romanians – 18% of the total population – lived abroad in 2019 (Table 1). In relative terms, after 2000 Latvia and Lithuania lost more than 10% of their respective populations via net migration, and Bulgaria and Romania lost 5%–10% (Figure 3). Most of these migrants headed for the large labour markets in Western Europe, namely Germany and the UK. Few CEE migrants live in other CEE countries; for example, only 28,000 Bulgarians out of 870,000 emigres (3%) live in another CEE country.

Table 1: Summary statistics: Countries

	Population (January 2019)	Number of citizens living in another EU country (2018)	As % of the total population	Citizens of other EU countries living in the country	As % of the country's population
Bulgaria	7,000,039	872,326	12.5%	13,696	0.2%
Czechia	10,649,800	163,990	1.5%	232,511	2.2%
Estonia	1,324,820	87,222	6.6%	20,891	1.6%
Croatia	4,076,246	523,886	12.9%	17,995	0.4%
Latvia	1,919,968	193,457	10.1%	6,433	0.3%
Lithuania	2,794,184	390,193	14.0%	7,483	0.3%
Hungary	9,772,756	446,587	4.6%	74,266	0.8%
Poland	37,972,812	2,475,906	6.5%	31,644	0.1%
Romania	19,414,458	3,533,186	18.2%	60,265	0.3%
Slovenia	2,080,908	68,008	3.3%	20,700	1.0%
Slovakia	5,450,421	342,682	6.3%	58,308	1.1%
Memo					
CEE	102,456,412	9,097,443	8.9%	544,192	0.5%
EU-28	513,471,676	17,608,436	3.4%	17,859,499	3.5%
Germany	83,019,213	889,484	1.1%	4,383,694	5.3%
France	67,012,883	776,308	1.2%	1,604,398	2.4%
Netherlands	17,282,163	563,396	3.3%	567,724	3.3%
Austria	8.,858,775	223,678	2.5%	730,209	8.2%
Portugal	10,276,617	1,195,934	11.6%	158,915	1.5%
UK	66,647,112	856,862	1.3%	3,681,859	5.5%

Source: Eurostat.

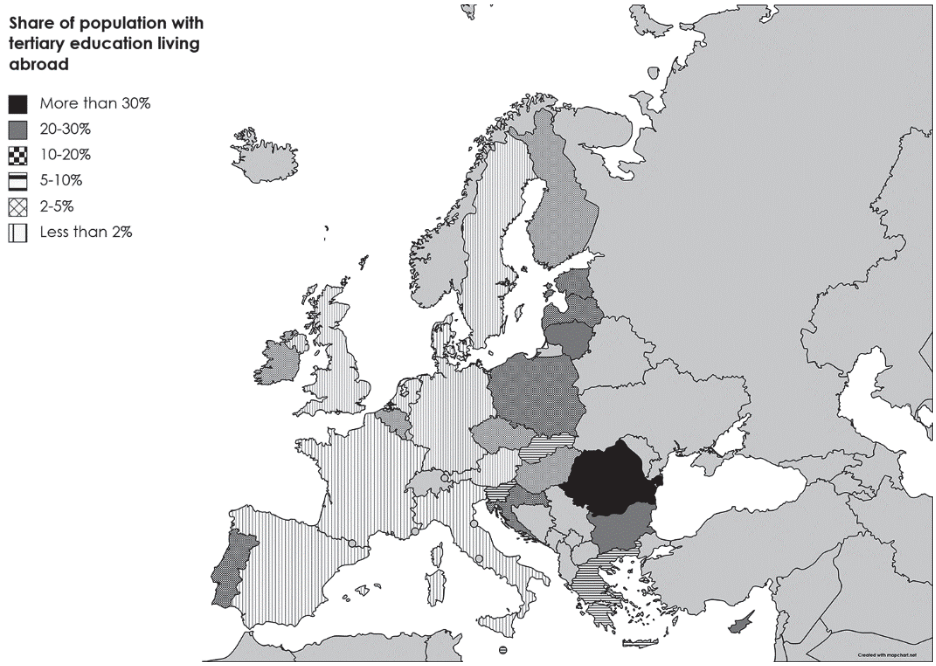
Figure 3: Net migration in the EU, 2000–2019



Source: Eurostat database, Annual Regional Database, author's calculation.

In order to examine Kaczmarczyk and Okolski's argument that migration in the EU amounted to 'brain waste', we also calculated the share of the each country's population with tertiary education that lives in another EU country. This number should serve as a proxy for the brain drain. Indeed, more than 30% of Romanians with a college degree left the country, and 20% of college-educated citizens of many other CEE countries lived abroad in 2017. To put these numbers in context, in most Western European countries except Portugal the share is typically less than 2% (Figure 4).

Figure 4: Share of tertiary education diploma holders living abroad, 2018



Source: Eurostat database, Annual Regional Database, author's calculation.

6. NUTS2 REGIONAL-LEVEL STYLIZED FACTS

Eurostat's regional data combined with the Commission's Regional database allows a more granular analysis of migration flows at the NUTS2 regional level. We use detailed data on 268 regions over a 14-year data span (2005–2018), with almost 3,500 observations. Table 2 summarizes the main statistical information in our dataset.

The largest NUTS2 region (Ile de France, with 12 million inhabitants) is 100 times more populous than the smallest (Valle d'Aosta in Italy, with 60,000 inhabitants). The per capita purchasing power in the wealthiest region – Inner London West – has a GDP more than 20 times higher than the poorest region – North-West Bulgaria. Inner London West, which includes the City of London, is in many respects an outlier. Its GDP per head is between 550%–600% of the EU average, and the hourly wage is similarly inflated. Productivity, measured by GDP per active worker, averaged more than €300,000 in Inner London West during 2005–

2018, more than twice as much as in the next region (Luxembourg). Except for South West Bulgaria, which includes the capital Sofia, the Bulgarian regions averaged 50-times lower productivity than Inner London-West, suggesting that we need to use London's numbers with caution.[10]

Table 2: Summary statistics: NUTS2 regions

3,752 observations (3,484 observations for migration data)	Minimum	Maximum	Average	Coef. of variation
GDP per capita (in PPS)	5,900	190,500*	26,446	48.2
GDP per capita (in % of EU)	23.9	628.0*	97.2	47.5
GDP per capita Annual Growth	-16.4	47.8	2.2	197.3
Productivity (GDP per active)	12,488	303,844*	51,435	42.2
Investment (% of GDP)	7.4	66.4	21.2	24.1
Unemployment rate (%)	1.3	37.0	8.6	62.4
LT Unemployment rate (%)	0.3	22.9	3.8	92.3
Employment Rate (%)	42.1	83.2	69.6	10.9
Wage (euro/hour)	1.1	44.1*	16.9	50.6
Population (thousand)	67.6	12,210.5	1,868.4	81.7
Active Population (%)	23.0	68.1	49.4	9.8
Median Age	31.4	50.7	41.7	7.5
Tertiary Education (%)	6.8	74.7*	27.3	35.7
Secondary Education	10.5	79.6	47.2	30.7
Rate of population change (‰)	-11.8	34.8	0.3	1116.1
Net migration (‰)	-25.2	55.2	2.7	193.5

Source: Eurostat database, Annual Regional Database.

Note: Minimum and Maximum are calculated from yearly regional observations.

* numbers for Inner London West.

Labour markets vary significantly across the EU regions. There is an apparent rift between north (Scandinavia, British Isles) and south (Italy, Spain, Greece) Europe concerning employment activity: while 79% of eligible people are in the labour force in Sweden (77% in the Netherlands, 75% in the UK), only 61%–62% work in Greece, Italy, and Spain. Central and Eastern European countries mostly fall somewhere between these poles. Only Czechia and Estonia record 'northern'

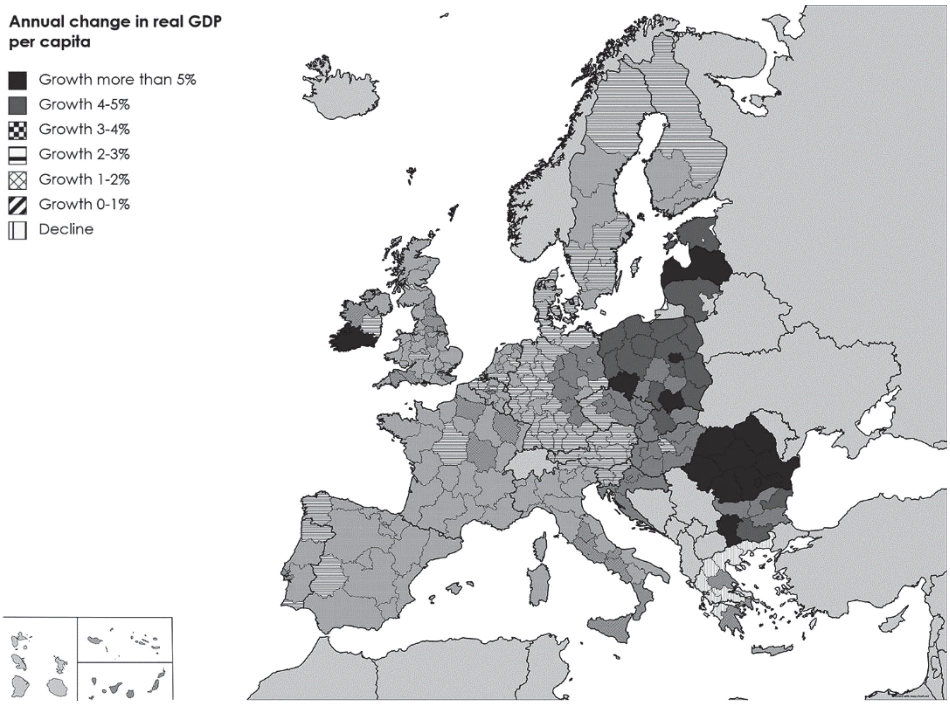
10 To control for the Inner London exceptionalism, we ran all the regressions with and without Inner London data, but the differences were minimal.

employment rates above 70%, while Hungary is firmly in the 'southern' camp with a 62% employment rate.

About one-quarter of all inhabitants in the 268 regions had tertiary education, with London regions boasting a share above 50%. The percentage was as low as 11% in Romania and the southern Italy regions. The lowest tertiary education share of 6.8% was recorded in the Czech South-West region in 2008 (it had recovered to 14% by 2018).

The growth rate in 2006–2018, measured by a change in real GDP per capita, averaged 2.2% across all the regions. However, the variance was immense: while in 10 Greek regions GDP per head in purchasing power parity was lower in 2018 than in 2006, the 8 Romanian regions grew at an annual rate of 7.4%. Figure 5 suggests that growth was concentrated in eastern (except for Greece) and northern regions and was notably weaker in western and southern regions.

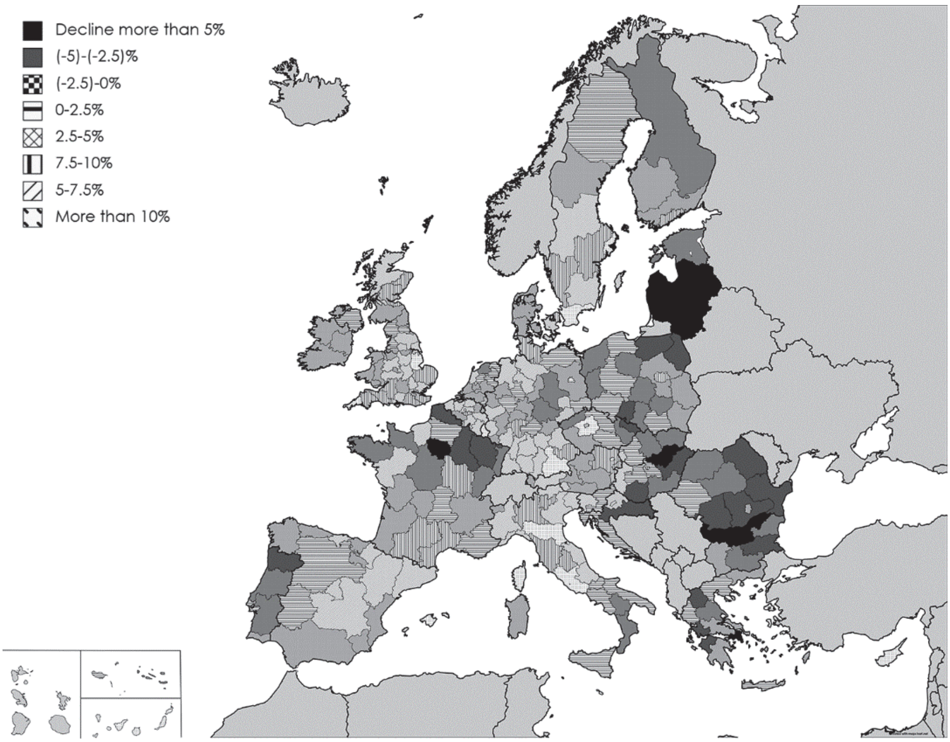
Figure 5: Real GDP per capita growth across the EU regions, 2006–2018



Source: Eurostat database, Annual Regional Database, author's calculation.

It is worth noticing that the total rate of population change is the most heterogeneous variable in the sample with the highest coefficient of variation. There were 18 instances when more than 10‰ citizens of a given region left in a given year, 7 of them in Latvia or Lithuania, between 2008 and 2011. During 2006–2018, Latvia and Lithuania lost more than 10% of their respective populations through net migration (Figure 6). Bulgaria lost 2.8% of its population via migration, but its most affected region – North-West – lost 8.5%. The worst affected region in 'old' EU member countries – Greece's Attiki – lost 6% of its population due to emigration in the 2006–2018 period. Lithuania suffered from the largest emigration rate in a single year: 25‰ citizens of the Baltic republic left in 2010.

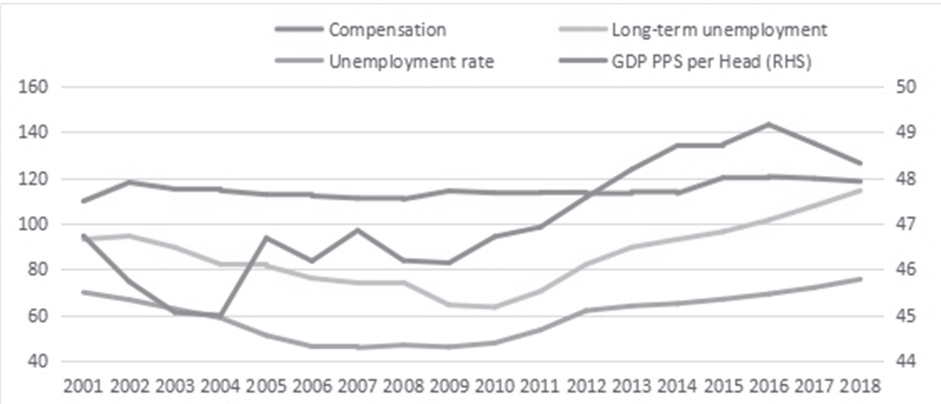
Figure 6: Migration in NUTS2 Regions, 2006–2018



Source: Eurostat database, Annual Regional Database, author's calculation.

Following Huber (2012), we further analysed regional disparities by calculating coefficients of variation for leading economic indicators: regional per capita income, compensation, and unemployment rates (standard and long-term).[11] By far the largest disparities are in wages (calculated as compensation per hour), followed by long-term unemployment and unemployment rate. GDP per head (in purchasing power) is the most consistent indicator across the 268 regions. Our analysis suggests that compensation disparities remained broadly constant in 2001–2018, with a small increase in 2015 negating earlier gradual decreases. Variance in GDP per capita steadily increased after the 2009 financial crisis until 2016 but declined thereafter. Labour market variance indicators decreased in 2001–2010; they re-emerged after the great financial crisis and kept widening even after (Figure 7). Both unemployment indicators were higher in 2015 than in 2001, showing a high hysteresis effect in unemployment rates, especially in southern member states.

Figure 7: Regional disparities: Coefficient of variation 2001–2015



Source: Eurostat database, Annual Regional Database, author's calculation.

7. RESULTS

The results section presents estimates of convergence in GDP per capita and unemployment rates across 268 regions without and with migration as an exogenous variable, to detect the impact of the latter. We also show the results of

11 We should stress that the variance coefficient measures only the sample's variance: it does not determine whether individual regions were converging or diverging.

the basic migration equations, where migration is the endogenous variable. We report the estimators and the probabilities for each estimator in brackets using the standard 10%, 5%, and 1% significance levels.

GDP Convergence Regression

Table 3 presents the results of four different models estimating real GDP per capita (y_t) convergence in the EU NUTS2 regions. The first specification estimates a simple convergence equation with the lagged variable of real GDP per capita (y_{t-1}) and the investment rate (inv_t). At 0.8 the estimated beta convergence coefficient is lower than in Huber (2012), suggesting a faster convergence, most likely owing to our broader and more heterogeneous data sample. Investments (inv_t) exhibit an expected positive and highly significant impact on GDP per capita. The Arellano-Bond tests for first- and second-order serial correlation in errors reject the autocorrelation hypothesis.

We examine the impact of migration on real GDP convergence in three regression models. We expand the convergence model by adding the exogenous variables net migration ($netmig_t$) and augmented natural population growth ($popgrowth_t$).[12] We also use the share of the population with tertiary education ($education_t$) as a measure of the education level in a given region.[13]

12 Following Fidrmuc (2019), we add 6 percentage points to the raw natural rate of population growth to offset the negative population growth in several regions during the 2000s.

13 We tested other specifications, with added long-term unemployment rate or wages, similar to Huber (2012). However, our estimates exhibited instability caused by strong correlation between unemployment and long-term unemployment rates and between education and wages, respectively. For these reasons, we report only short specifications that should be more robust.

Table 3: Regressions for real GDP per capita, dependent variable $ln(y_t)$

	Without migration	With net migration	Only immigration regions	Only emigration regions
$\ln(y_{t-1})$	0.891***	0.821***	0.777***	0.811***
	(0.007)	(0.008)	(0.012)	(0.017)
$\ln(inv_t)$	0.009***	0.036***	0.052***	0.019*
	(0.005)	(0.005)	(0.006)	(0.010)
$netmig_t$		0.011***	0.011***	0.019***
		(0.002)	(0.002)	(0.008)
$\ln(popgrowth)_t$		−0.487***	−0.455***	−0.567 ***
		(0.046)	(0.054)	(0.105)
$\ln(education)_t$		0.049***	0.060***	0.066***
		(0.007)	(0.008)	(0.018)
Adj. R^2	0.991	0.991	0.989	0.991
Obs.	3484	3484	2458	1026
Schwarz criterion	−3.04	−3.12	−3.09	−2.57
Durbin Watson	1.853	1.877	1.893	2.007

Source: Author

Note:* p<0.1, ** p<0.05, *** p<0.01

The first of these expanded regressions uses data from all 268 NUTS2 regions, while the remaining two split the data into two samples, for regions gaining or losing net migrants. The 'immigration' set consists of 1,846 observations, while the 'emigration' dataset is smaller at 834 observations. The lagged (y_{t-1}) variable effect is significant at the 1% level and in the 0.6–0.7 range in all three models. Similarly, investment and education effects provide consistent estimators with high significance in all models. Education level appears to increase real GDP per capita by 0.1%–0.2% for each percentage point of tertiary-educated inhabitants of a region. Natural population growth, not surprisingly, reduces real GDP per capita in all models, but its effect appears much stronger in the 'emigration' regions.

Most importantly, the impact of net migration on real GDP per capita is significant and positive in all three models. A 1 percentage point increase in net emigration should increase GDP per capita by 0.04%–0.05%, and by even 0.14% in the emigration regions. These estimates are higher than in Huber (2012), primarily due to a longer time period and larger sample. We include the years after the 2009 recession that increased volatility in GDP growth and intensified migration flows. Also, our sample comprises more extreme regions, such as Bulgarian and Romanian regions that lost 10% of their population between 2006 and 2015, or London, which gained substantial immigration flows throughout the decade.

Unemployment Rate Convergence Regression

The effects of migration on the unemployment rate are presented in Table 4. We show four models, a straightforward convergence model based on wage differentials and three models incorporating net migration flows. First, we estimated a simple convergence model whereby the unemployment rate (un_t) was estimated on the lagged variable (un_{t-1}) and the wage rate per hour ($wage_t$). Our results confirm convergence in unemployment rates among the EU regions with the lagged coefficient at 0.78. Wages contribute to higher unemployment, as expected.

The three migration regressions – one on the full sample and one for each 'immigrant' and 'emigrant' region – extend the simple model by adding the variables net migration ($netmig_t$) and augmented natural population growth ($popgrowth_t$).[14] The models provide consistent estimates of the unemployment hysteresis, with the lagged coefficient at 0.6–0.7. The wage coefficient is similarly consistent and significant across the models at approximately 0.1. The wage effect is significantly stronger in 'emigration' regions, suggesting a stronger pull effect of higher wages in richer regions. The population growth effect is insignificant in all models, which is in line with previous research (Huber, 2012). Most importantly, the impact of net migration on the unemployment rate is consistently negative across all models, reducing unemployment by approximately 0.1 percentage points per each percentage point of net migration.

14 A strong correlation between education and wages excluded the education variable in this model.

Table 4: Regressions for unemployment rates, dependent variable *ln(un$_t$)*

	Without migration	With net migration	Only immigration regions	Only emigration regions
ln(un$_{t-1}$)	0.780***	0.697***	0.696***	0.625***
	(0.012)	(0.015)	(0.018)	(0.028)
ln(wage$_t$)	0.096***	0.128***	0.094*	0.276***
	(0.037)	(0.039)	(0.053)	(0.061)
netmigrat		–0.112**	–0.102***	–0.191***
		(0.009)	(0.051)	(0.036)
ln(pop. growth)		–0.007	–0.001	0.005
		(0.006)	(0.008)	(0.009)
Adj. R^2	0.891	0.897	0.873	0.896
Obs.	3484	3484	2458	1026
Schwarz criterion	0.01	–0.05	0.13	0.24
Durbin Watson	1.609	1.580	1.714	1.558

Source: Author.

Note: * p<0.1, ** p<0.05, *** p<0.01

Migration Regression

Reversing the causality, we finally regress migration flows on economic factors. We simplify Marques's (2010) model in order to avoid multicollinearity and endogeneity issues and regress migration flows on only four exogenous variables: unemployment rate (un_t), wage level lagged one year ($wage_{t-1}$), and the lagged migration variable.

We begin with a simple migration model, with unemployment and wages as only exogenous variables. Our estimates in Table 5 confirm Marques's finding that the unemployment rate impacts migration negatively. The wage effect is weaker and significant only at the 10% level. The full migration model exhibits strong path-dependency, with the coefficient of $migration_{t-1}$ higher than 0.5 and highly significant. Higher unemployment rates are uniformly negative for migration, even in the CEE region (672 observations), and the wage effect is statistically insignificant when the sample is split into two subsets.

Our results confirm that it is statistically difficult to find robust and consistent estimators of the role economic factors play in migration flows within the European Union. While our qualitative analysis clearly shows that it flows from east to west and north, the regression results with respect to wages are ambiguous.

Table 5: Migration equations (dependent variable=net migration)

	Simple	With net migration	Only non-CEE regions	Only CEE regions
ln(un)	–0.545 ***	–0.183 ***	–0.218***	–0.070***
	(0.022)	(0.022)	(0.027)	(0.027)
ln($wage_{t-1}$)	–0.111*	–0.057	–0.057	–0.024
	(0.057)	(0.055)	(0.086)	(0.047)
$Migration_{t-1}$		0.550***	0.544***	0.496***
		(0.016)	(0.018)	(0.036)
Adj. R^2	0.500	0.680	0.590	0.770
Obs.	3484	3216	2544	672
Schwarz criterion	1.41	1.09	1.22	0.01
Durbin Watson	0.927	2.010	2.022	1.752

Source: Author

Note: * p<0.1, ** p<0.05, *** p<0.01

8. CONCLUSIONS

This paper investigates the current migration trends in the European Union, paying special attention to the uneven changes in the population of the 11 new member states that have become members of the EU since 2004. We illustrated momentous changes in some of these countries, namely the poorer states in the south and Eastern Europe. The scope of depopulation in Bulgaria, Romania, and the Baltic countries has no parallel in peacetime and may undermine these countries' long-term growth, and even viability. A loss of population near 10% in the decade up to 2015 reflects a low natural rate of population change, and our analysis suggests that outmigration from Eastern Europe contributed to the phenomenum.

We analyse the main driving forces behind migration and the extent to which migration contributes to convergence in incomes and unemployment rates. Our sample contains data on 286 NUTS EU regions for the years 2005–2015. We were able to significantly extend the data sets used in previous research by adding volatile data from the 2008–2009 crisis and the post-crisis years 2010–2015. We were also able to include data on regions in Bulgaria, Croatia, Romania, and the United Kingdom. Our econometric analysis suggests that migration positively impacts convergence in GDP per capita.

We specified three models using the sample of almost 3,000 observations and two subsets, and we estimated coefficients of net migration using the fixed-effects method. They suggest that each percentage point in migration increases GDP per capita by 0.04%–0.1%. These estimates are higher than in most of the literature from the 2000s, mostly owing to a longer time period and larger sample.

We also estimated the effects of migration on unemployment rates in the EU regions. Our estimates affirm that unemployment is strongly path-dependent and that a 1 percentage point increase in wages typically increases the unemployment rate by roughly 0.2. Most importantly, our regressions suggest that the impact of net migration on the unemployment rate is consistently negative, while previous studies were mostly unable to find a statistically significant effect. Our more robust estimates are due to an extended sample, with more regions and more observations.

While we were able to determine the effects of net migration on the main economic variables – GDP per capita and unemployment rate – we were less successful in estimating the inverse relation. The estimated effects of the unemployment rate and wages on migration are either insignificant or counterintuitive. As previous researchers have noted, migration decisions are complex, rooted in characteristics that are only marginally captured by macroeconomic variables.

APPENDIX

Table A1: National statistics

	Minimum	Maximum	Average	Coef. of variation
GDP per capita (in PPS)	6,138	171,006*	25,625	46.2
GDP per capita (in % of EU)	24.8	599.7*	97.3	45.8
GDP per capita Annual Growth	–16.4	35.9	2.1	219.0
Productivity (GDP per active)	12,488	303,844*	51,435	42.2
Investment (% of GDP)	8.4	65.8	21.8	24.7
Unemployment rate (%)	1.9	37.0	8.9	59.3
LT Unemployment rate (%)	0.4	22.9	3.9	87.7
Employment Rate (%)	42.1	83.2	69.6	10.9
Wage (euro/hour)	1.1	105.5*	16.6	57.1
Population	123,598	12,106,455	1,863,246	81.6
Active Population (%)	23.0	68.1	49.4	9.8
Median Age	31.4	50.1	41.3	7.3
Tertiary Education (%)	6.8	69.8*	26.1	35.7
Secondary Education	10.5	79.6	47.2	30.7
Rate of population change (‰)	–11.8	34.8	0.3	1116.1
Net migration (‰)	–25.2	55.2	2.7	196.5

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THE EFFECT OF THE COVID-19 PANDEMIC ON CONSUMER SAVINGS AND RETAIL SALES: EVIDENCE FROM A POST-COMMUNIST TRANSITION ECONOMY

ABSTRACT: *When coupled with strong external shocks such as COVID-19, the high levels of uncertainty that characterise fragile economies can have a strong impact on household consumption and saving behaviour. This paper analyses household consumption and saving behaviour in conjunction with COVID-19 in the context of a post-communist economy. Models and intervention analysis are used to identify the effect of catastrophic events such as the COVID-19 pandemic on two key macroeconomic measures for the Albanian economy. The findings show that the pandemic period caused a significant contraction of consumer spending and a significant increase in savings. Higher uncertainty appears to have been a key driver of such household behaviour. The effect on savings will endure in the long run, while retail trade is expected to recover. These findings call for a more astute use of fiscal and monetary policies to address the harmful emerging short-run effect of reduced household spending.*

KEY WORDS: *Albania, ARIMA modelling, catastrophic events, COVID-19 pandemic, intervention analysis*

JEL CLASSIFICATION: D14, E21, E27, E71

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1. INTRODUCTION

The COVID-19 pandemic has negatively affected the global economy through various channels. The effect of the pandemic on the economy can be expressed as a direct effect emerging from the voluntary reduction in transaction-based activities, especially contact-intensive ones (Lee et al. 2021). The result is a change in consumer behaviour, primarily a reduction in consumer spending on goods and services and an increase in savings. The pandemic has also produced an indirect effect on the economy following governmental decisions to enact lockdowns to reduce the spread of the virus. On the one hand, the lockdowns and restrictive measures contributed to a slow-down of economic activity, leading to shocks in both supply and demand, a decline in employment and production, and supply shortages. On the other hand, the COVID-19 pandemic has negatively affected aggregate demand through higher unemployment, lower income, and increased uncertainty, and consequently consumption (Baldwin and Mauro, 2020).

Previous research has used different frameworks to control for transmission channels and the heterogenous effects of pandemics in the economy. Maliszewska et al. (2020) utilise previous literature on the estimated effects of other epidemics such as SARS (Lee et al. 2004), avian influenza (Burns et al. 2006), and Ebola (Evans et al. 2014) to assemble a framework composed of four channels of effects, namely 1) the direct impact of a reduction in employment; 2) the increase in the cost of international transactions; 3) the sharp drop in travel; and 4) the decline in demand for services that require proximity between people.

As a result of the COVID-19 pandemic, the world has experienced a reduction in global trade volume and global gross domestic product (GDP). However, the estimated changes to key economic outcomes have differed depending on the country and the period under study (OECD 2020; IMF 2021; König and Winkler 2021; McKibbin and Fernando 2021). While COVID-19 has had negative economic impacts across the globe, given their limited capacity to cope with economic shocks, the effects have been more pronounced in developing or transition economies, making a severe impact on poverty rates more likely. Higher levels of economic uncertainty (which characterize fragile economies) coupled with strong external shocks (such as the one triggered by COVID-19) are expected to have a strong impact on household spending behaviour. This paper investigates the effects of the COVID-19 pandemic on two key macroeconomic measures in Albania – a post-communist transition economy, one of the lowest-income countries in Europe, and one of the economies most affected by the COVID-19 pandemic. Specifically, the paper investigates the effects of the recent COVID-19 pandemic on retail sales and consumer savings in Albania and the duration of these effects. The effect on retail sales is used as a proxy for the effect on the consumption level, while the effect on the savings rate is used to primarily measure the direct effect of the pandemic on consumer behaviour.

Several studies have concluded that the world faces the worst economic downturn experienced in the last decades. In the first few months of the pandemic, global GDP decreased by 5.2%, resulting in one of the deepest economic contractions since WWII (IMF 2021). The euro area economy also revealed strong signs of contraction: during the first and second quarters of 2020 GDP decreased by 3.7% and 11.7% respectively on a quarter-to-quarter basis (EC 2021). Maliszewska et al. (2020) estimated that global GDP for the year 2020 would decrease by 2%, with a 2.5% decrease for developing countries and 1.8% decrease for industrial countries. Global pandemic effects are expected to be more detrimental in countries with relatively higher trade integration and a larger tourism presence.

The scale of the economic crisis has also varied by country, depending on the pandemic's evolution and the lockdown measures taken by respective governments. In the United States, for example, approximately 20 million job losses were recorded by the second quarter of 2020 (Danielli et al. 2021). According to Lee et al. (2021), a strong increase in the unemployment rate was observed during the first part of 2020, with larger effects on women, minorities, the less educated, and the young, especially in the states with the highest prevalence of infection. The International Labour Organization (ILO) (2020) estimated a global reduction in working hours of slightly more than 10% during the second quarter of 2020 compared to the previous pre-crisis quarter. The effect was especially pronounced in reduced working hours, furloughs, and work-from-home arrangements (Cook and Grimshaw 2021; Bluedorn et al. 2021).

The COVID-19 pandemic and induced economic and social constraints have significantly impacted the confidence of both consumers and businesses. Despite the significant effects, there is a scarcity of comprehensive studies on the impact of the COVID-19 pandemic on consumer and business sentiment. A study by a large group of scholars (Altig et al. 2020), using several economic uncertainty indicators for the US and UK (implied stock market volatility, newspaper-based economic policy uncertainty, and other economic uncertainty indicators collected from social media) identifies large uncertainty jumps in reaction to the pandemic and its economic fallout. Teresiene et al. (2021), using the consumer confidence index (CCI), manufacturing purchasing manager's index, and services purchasing manager's index as dependent variables, find that the pandemic produced a rapid and robust effect in the short term, but that longer-term results depend on the region. The spread of the COVID-19 pandemic had a negative effect on the CCI in the US and China (Teresiene et al. 2021). Van der Wielen and Barrios (2021) document a substantial increase in people's unemployment concerns, to levels above those during the Great Recession. In addition, they observe a slowdown in consumption. The ensuing shift in sentiment was significantly deeper in countries hit hardest in economic terms (Van der Wielen and Barrios 2021).

This study, which was carried out during early 2021, contributes to the empirical literature measuring the economic effects of the COVID-19 pandemic in the context of a post-communist economy, Albania. The overall goal of this research is to gain a better understanding of the effects of the COVID-19 pandemic on key macroeconomic measures for the Albanian economy related to consumer behaviour. The specific objectives are twofold. First, the paper investigates the effect of the COVID-19 pandemic on consumer spending, measured through retail sales, and second, the paper investigates the effect of the COVID-19 pandemic on consumer savings in Albania.

We find a significant contraction of consumer spending and on the one hand and a significant increase of savings on the other. Higher uncertainty appears to have been a key driver of such household behaviour.

2. THE COVID-19 PANDEMIC AND THE ALBANIAN ECONOMY

Albania, a country with one of the lowest incomes in Europe, is also one of the countries most affected by the COVID-19 pandemic. The first signs of contraction of the Albanian economy due to the pandemic appeared in the first quarter of 2020, predominantly in the trade and tourism sectors. One of the most crucial economic challenges is the disruption of global value chains and international trade that emerges after hindered production and disturbances in demand or investments (Baldwin and Freeman, 2020), caused by the pandemic, and compounded by the high degree of globalisation. Albania has close trade and foreign direct investment (FDI) relations with European Union (EU) countries, especially Italy, which was also among the countries most effected by COVID-19. The contractionary demand in EU countries and the closure of borders affected the service-oriented sectors of the Albanian economy such as trade, transport, and tourism. The primary exported commodity group, cut-make-trim goods, was most affected by this disruption in trade. Diminished demand from the EU markets, default orders, and the creation of large stocks weakened and severely damaged the trade sector of the Albanian economy (Musabelliu 2020). Only three months into the pandemic, Albania experienced a decrease of 44.4% and 36.7% in the value of exports and imports respectively (United Nations 2020). The Bank of Albania (central bank of Albania) reported a 2.5% reduction in Albanian GDP in annual terms for the first quarter of 2020 (BoA, 2020), mainly due to the drop in investments and the trade of goods and services, and consequently final consumption.

Albania has also witnessed strong social effects of the COVID-19 pandemic, particularly in terms of the number of unemployed and poor people. The sudden fall in revenues and in overall demand caused liquidity shortages in many small and medium enterprises (SME), forcing them to lay off workers (United Nations, 2020).

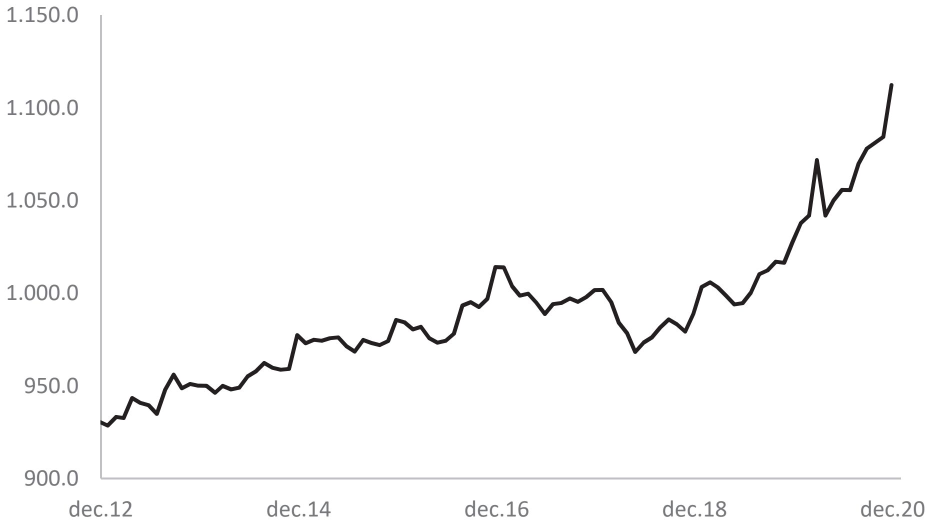
Data from the Albanian Labour Force Survey (INSTAT 2021b) show that the unemployment rate increased on a quarter-to-quarter basis by 0.2 percentage points in the first quarter of 2020 (from 11.2% to 11.4%) and by 0.5 percentage points in the second quarter (from 11.4% to 11.9%). Despite some small gains in the subsequent quarters, the unemployment rate increased again to 11.9% in the first quarter of 2021, demonstrating the imbalance in the labour market sector during winter periods. The effect of the pandemic was more pronounced for the young (the youth unemployment rate increased by 3.3 percentage points during the first quarter of 2021, from 20% to 23.3%). The official figures do not reveal the full effects of the pandemic in the labour market – many people were officially considered employed just because they continued to work on farms, disregarding the intensity of employment (e.g., working 10 hours versus 50 hours per week). Surveys by World Vision (2020) show that unemployment increased by 10%, while full-time employment decreased by 9%.

Albania suffered an increase in hidden unemployment due to the closure of the borders with Italy and Greece, which are key destinations for seasonal employment. Overall, a 14% decrease in occasional employment was recorded (ibid).

The pandemic negatively affected Albanian household salaries (both formal and informal wages) as well as remittances, which account for a significant part of consumption spending – remittances constitute around 9.5% of Albanian GDP (Musabelliu 2020). A World Bank report (2020) estimates that the Albanian economy experienced a 20% drop in remittances due to the economic crisis caused by COVID-19 pandemic, contributing further to decreased demand and consumption in Albania. The shock in the labour market was accompanied by a 1 percentage point increase in the poverty rate (WB2021a). According to World Bank estimates, the poverty rate would have been 1.8 percentage points higher if no response measures had been taken during 2020 (WB 2021b).

The negative economic and employment trends have been a major concern for Albanian households. A World Vision survey (2020) revealed that around 68% of the surveyed Albanian households expected their employment to be negatively affected by the COVID-19 pandemic, with people between the ages of 30 to 60 comprising the most affected group. These negative household expectations regarding employment prospects and income resulted in reduced household consumption (and increased savings) in order to prepare for future potential unemployment. Recent data show a substantial increase in savings deposits in Albania during 2020 (Figure 1). Unlike previous years, when the growth of deposits was driven by foreign currency deposits while savings/deposits in local currency (Albanian lek) were not as attractive (due to falling interest rates since 2012), savings during the pandemic have been oriented towards the Albanian lek. Domestic currency deposits increased by 11% in 2020. This is an indirect indication that domestic savings drove the increase in deposits (Monitor 2021).

Figure 1: Savings in billion Albanian Lek for the period December 2012 – December 2020.



Source: Bank of Albania (2021)

Previous reports on or analysis of the impact of COVID-19 on the Albanian economy have mostly relied on basic descriptive statistical analysis. There is a lack of rigorous statistical analysis of the impact of the COVID-19 pandemic on savings and consumer behaviour, and this paper contributes to filling this gap.

3. DATA AND EMPIRICAL PROCEDURE

The data used in this study consists of monthly observations of retail sales and the savings rate in Albania. Data were obtained from the Albanian Institute of Statistics (INSTAT 2021a). Both savings and retail trade are important components of the Albanian economy: savings make up 11% of GDP (World Bank 2021) and retail trade 16.8% of GDP (INSTAT 2021a). These are the highest-frequency data available for Albania on macroeconomic measures relevant to the study of the effect of the COVD-19 pandemic. However, even these data are published with a delay of at least three months.

We employ intervention analysis (Enders 1995, pp. 240–247) to identify a change in the mean of a stationary time series. The impacts of a stochastic catastrophic event like the recent COVID-19 pandemic can be measured ex post by using dummy variables. Let D_t^τ be a 'pulse' dummy variable defined as

$$D_t^\tau = \begin{cases} 0, & t \neq \tau \\ 1, & t = \tau \end{cases} \tag{1}$$

where τ represents the time of the catastrophic event, and let S_t^τ be a 'step' dummy variable defined as

$$S_t^\tau = \begin{cases} 0, & t < \tau \\ 1, & t \geq \tau \end{cases} \tag{2}$$

The pulse function (1) is used to test for a short-run impact associated with the COVID-19 pandemic. The step function (2) is used to test for a structural adjustment due to long-term effects of the pandemic on consumer behaviour in Albania.

We use an ARMA (*p*,*q*) process to model the dynamic properties of the retail sales and savings rate series:

$$y_t = f(y_{t-p}, \varepsilon_{t-q}, D_t^\tau, S_t^\tau) \tag{3}$$

where p and q represent respectively the order of the autoregressive and moving average process and ε_t is a white noise error with mean zero and constant variance.

As an example, for an AR (1) specification (3) becomes

$$y_t = \alpha_0 + \alpha_1 y_{t-1} + \alpha_2 D_t^\tau + \alpha_3 S_t^\tau + \varepsilon_t \tag{4}$$

where $0 < \alpha_1 < 1$. In the case of retail sales, the reduction in retail sales following the pandemic suggests that α_2 is negative, while the gradual recovery in retail sales following the initial shock suggests that α_3 is positive. The initial overall effect of the pandemic on retail sales is then calculated as $(\alpha_2 + \alpha_3)$. The change in long-run equilibrium is *LR* = $\alpha_3 / (1 - \alpha_1)$ and is calculated as the difference between the long-run mean after the pandemic, $(\alpha_0 + \alpha_3)/(1 - \alpha_1)$, and the long-run mean before the pandemic, $\alpha_0 / (1 - \alpha_1)$.

The dynamic effects of the pandemic can be obtained from the impulse response function obtained by applying the lag operator to (4) (Enders1995) and is given by

$$y_t = \alpha_0 / (1 - \alpha_1) + \alpha_2 \sum_{i=0}^{\infty} \alpha_1^i D_{t-i}^{\tau} + \alpha_3 \sum_{i=0}^{\infty} \alpha_1^i S_{t-i}^{\tau} + \sum_{i=0}^{\infty} \alpha_1^i \varepsilon_{t-i} \tag{5}$$

Equation (5) enables us to trace out the dynamic impacts of pandemic on the time path of retail sales by differentiating y_t with respect to D_t^{τ} and S_t^{τ} at different point in times dy_{t+i} / dD_t^{τ} and dy_{t+i} / dS_t^{τ} for all *i*> 0.

It is important to note that (4) and (5) can be modified for different AR(*p*) or MA(*q*) processes. Additionally, before identifying and estimating the correct specification of (4) for each series, both series are first tested for stationarity using the augmented Dickey-Fuller (ADF) test (Dickey and Fuller 1979) and the Phillips and Perron (1988) unit root test.

4. EMPIRICAL RESULTS

Stationarity Tests

Results of the unit root testing are presented in Table 1. Both the Dickey-Fuller (1979) and the Phillips and Perron (1988) unit root tests reported in the top panel of Table 1 indicate that the savings series is nonstationary. This is true whether one assumes that the savings series has no drift (zero mean), has a drift (single mean), or has a trend. Therefore, first-differencing was performed on the savings series. The first-differenced series was tested again for stationarity. Both the Dickey-Fuller (1979) and the Phillips and Perron (1988) unit root tests reported in the bottom panel of Table 1 indicate that the differenced savings series is stationary.

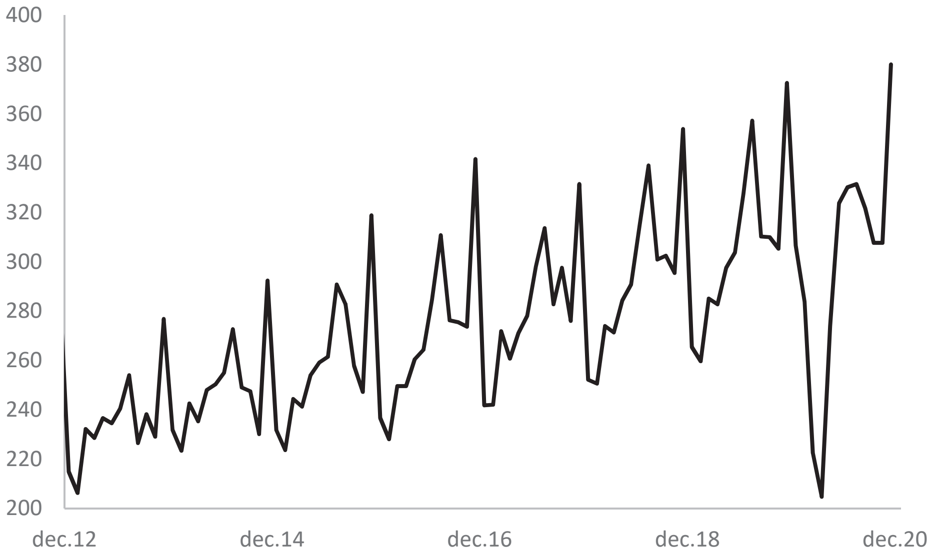
Table 1: Stationarity Tests for the presence of unit roots in savings and retail sales

	Savings				Retail Sales			
	Series in levels							
	Phillips-Perron Unit Root Test							
Type	Rho	Pr < Rho*	Tau	Pr < Tau*				
Zero Mean	0.6211	0.8355	4.2853	1				
Single Mean	-1.6211	0.8212	-2.2421	0.192				
Trend	-3.1252	0.9305	-1.6291	0.7781				
	Augmented Dickey-Fuller Unit Root Tests							
Type	Rho	Pr < Rho*	Tau	Pr < Tau*				
Zero Mean	0.6234	0.8357	4.8472	0.9999				
Single Mean	-1.5834	0.8252	-2.3695	0.1519				
Trend	-2.8482	0.942	-1.612	0.7845				
	First-differenced				*Twelfth-differenced*			
	Phillips-Perron Unit Root Test				Phillips-Perron Unit Root Test			
Type	Rho	Pr < Rho*	Tau	Pr < Tau*	Rho	Pr < Rho*	Tau	Pr < Tau*
Zero Mean	-123.9	<.0001	-9.5853	<.0001	-45.921	<.0001	-5.1443	<.0001
Single Mean	-146.33	0.0013	-10.977	<.0001	-66.996	0.0015	-6.2889	<.0001
Trend	-149.71	0.0006	-11.108	<.0001	-66.986	0.0006	-6.2754	<.0001
	Augmented Dickey-Fuller Unit Root Tests				Augmented Dickey-Fuller Unit Root Tests			
Type	Rho	Pr < Rho*	Tau	Pr < Tau*	Rho	Pr < Rho*	Tau	Pr < Tau*
Zero Mean	-122.5	<.0001	-9.5576	<.0001	-54.293	<.0001	-5.5287	<.0001
Single Mean	-144.4	<.0001	-10.952	<.0001	-73.036	0.0015	-6.5173	<.0001
Trend	-148.04	<.0001	-11.088	<.0001	-73.053	0.0006	-6.5056	<.0001

Note: *These are p-values.

The retail sales series, depicted in Figure 2, have a well-defined seasonal pattern. Retail sales are low in the first few months of the year, increase during the months of June, July, and August, fall during the months of September, October, and November, and rise in a pronounced spike in the month of December. This pattern repeats regularly almost every year. A twelfth difference was therefore applied to the retail sales series to eliminate this pronounced seasonal pattern. The twelfth-differenced series was then tested for stationarity. Both the Dickey-Fuller (1979) and the Phillips and Perron (1988) unit root tests reported in the bottom panel of Table 1 indicate that the twelfth-differenced retail sales series is stationary.

Figure 2: Retail Sales Index (March 2005=100) for the period December 2012 – December 2020.



Intervention Analysis Results

Table 2 reports the results of the intervention analysis. The Box and Jenkins (1970) approach is used to identify the appropriate ARMA (p, q) specification for both savings and retail sales. The data prior to the intervention event (the COVID-19 pandemic) are used to identify the appropriate ARMA (p, q) specification (Enders 1995). The World Health Organization (WHO 2020) declared the COVID-19 infection a pandemic on 11 March 2020. Therefore, data prior to March 2020 are used to identify the appropriate ARMA (p, q) specification. Once the appropriate ARMA (p, q) specification is identified, the complete data are used to estimate (3) by including the pulse and the step dummy variables in the model.

Table 2: Intervention analysis results for savings and retail sales

Variable	Savings	Retail Sales
Constant	3199***	-
y_{t-1}	0.212***	0.642***
y_{t-3}	-	0.150***
y_{t-12}	0.351***	-
D_t^{τ}	34755***	-89.798***
S_t^{τ}	7459**	8.987***
ε_{t-1}		0.981***
ε_{t-12}		0.525***
AIC	3488.6	1910.3
BIC	3504.2	1931.5
Log Likelihood	-1,739.30	-949.2
Short-run and long-run effects of the pandemic		
Short-run effects	42,214 (4.2%)	-80.81 (-26.6%)
Long-run effects	17,096 (1.7%)	43.28 (14.2%)

Note: *, **, and *** denote significance level respectively at 10%, 5%, and 1%.

An AR (1, 12) specification is fitted to the first-differenced series of savings. As noted earlier, stationarity tests indicate that the twelfth-differenced series of retail sales is stationary. An AR (1, 3, 12) specification provides the best fit for this series. However, the autocorrelation function (ACF) indicates a very high correlation at the first lag and a slowly decaying correlation function. Additionally, autocorrelation tests of residuals indicate the presence of remaining autocorrelation.[1] Based on this, the first difference is applied to the twelfth-differenced series of retail sales. The twice-differenced series is then used to identify the appropriate ARMA (*p*, *q*) specification. An ARMA specification with AR (1, 3) and MA (1, 12) specification provides the best fit for this series. Akaike

[1] The results for this specification and diagnostic test results are not reported here but are available upon request.

information criterion (AIC) and Bayesian information criterion (BIC) indicate that this specification is superior to the AR (1, 3, 12) specification of the twelfth-differenced series of retail sales. Additionally, the residuals resemble white noise with no presence of remaining autocorrelation.

The results of the AR (1, 12) specification for the first-differenced series of savings and the ARMA specification with AR (1, 3) and MA (1, 12) specification for the retail sales series are presented in Table 2. Table 2 also reports the AIC, BIC, and log likelihood values for both series. The results of the null hypothesis that residuals are white noise are presented in Table 3 and indicate no autocorrelation present in the residuals.

Table 3: Results of the tests of the null hypothesis that residuals for savings and retail sales are white noise

	Savings			Retail Sales		
Lag length	Chi-Square	DF	Pr > ChiSq	Chi-Square	DF	Pr > ChiSq
6	3.95	4	0.4123	0.89	2	0.6422
12	12.1	10	0.2786	1.53	8	0.9923
18	20.62	16	0.1936	3.95	14	0.9957
24	31.78	22	0.0812	7.91	20	0.9925
30	33.66	28	0.2122	8.97	26	0.9992

Note: A value of "Pr>Chisq" greater than a chosen significance levels (i.e., 5%) indicates that residuals are white noise.

Table 2 also reports the short-run and long-run effects of the pandemic on savings and retail sales. The short-run effect of the pandemic on savings is an increase of 42,214 million Albanian Lek (ALL),[2] a 4.2% increase on the average of the two-year period prior to the pandemic (March 2018 – February 2020). The long-run effect of the pandemic on savings is an increase of 17,096 million Albanian Lek (ALL) or a 1.7% increase on the average of the two-year period prior to the pandemic.

2 As of 21 May 2021, the USD/ALL exchange rate was 101.

The short-run effect of the pandemic on retail sales is a decrease of 80.81 points (the retail sales series is an index with the March 2015 value of the index equalling 100), a 26.6% decrease on the average of the two-year period prior to the pandemic (March 2018 – February 2020). The long-run effect of the pandemic on retail sales is an increase of 43.28 points or a 14.2% increase on the average of the two-year period prior to the pandemic.

Figures 3 and 4 present the impulse response functions for savings and retail sales. The left vertical axis in Figures 3 and 4 depicts the percentage change, while the right vertical axis in both figures depicts the actual units for each series. Figure 3 shows an initial increase in savings of 42,214 million ALL (4.2%) in March 2020 and a gradual decrease in the rate of increase in savings to a long-run effect of 17,096 million ALL (1.7%). Figure 3 also shows that the long-run effect is mostly realised by September 2020. Figure 4 shows an initial decrease in retail sales by 80.81 points (26.6%) in March 2020 and a gradual increase in retail sales to the long-run effect of 43.28 points (14.2%). About 90% of the long-run effect is realised by May 2021.

Figure 3. Impulse response Function Following the COVID-19 Pandemic for Savings in Albania

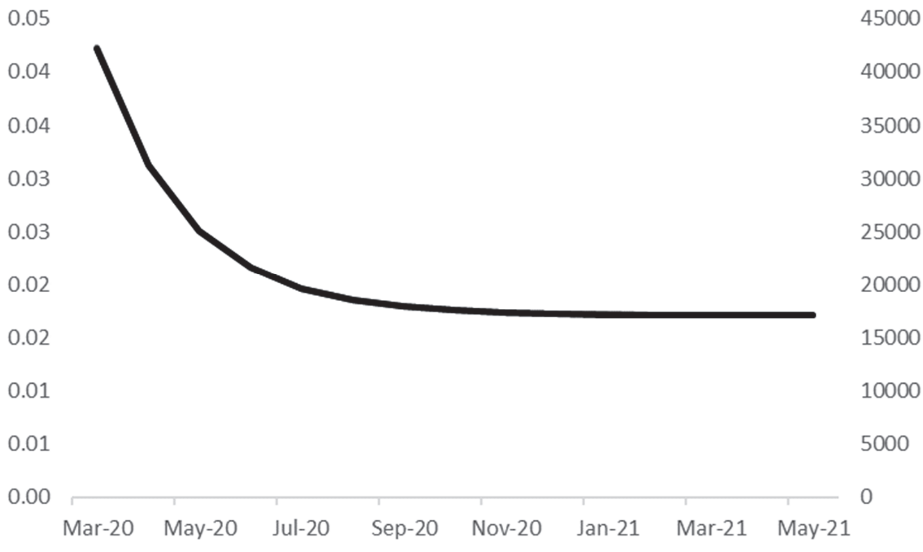
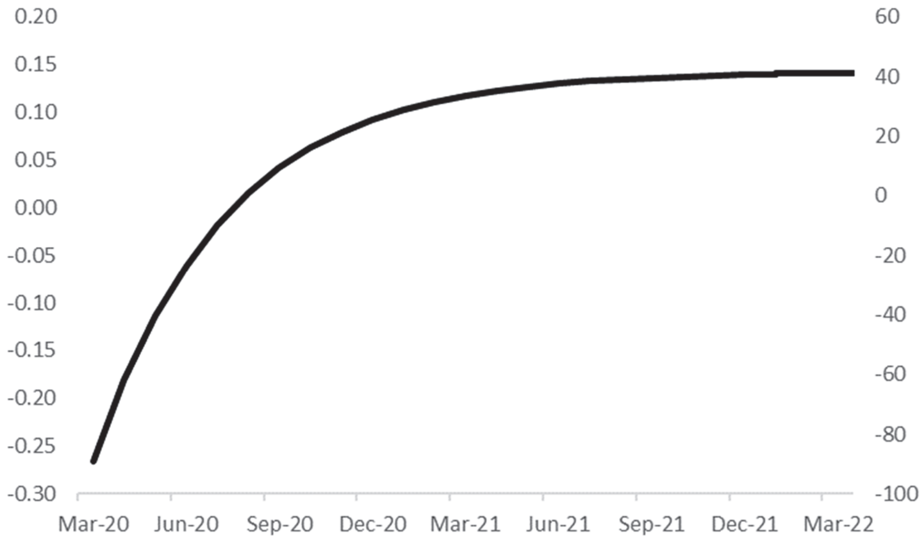


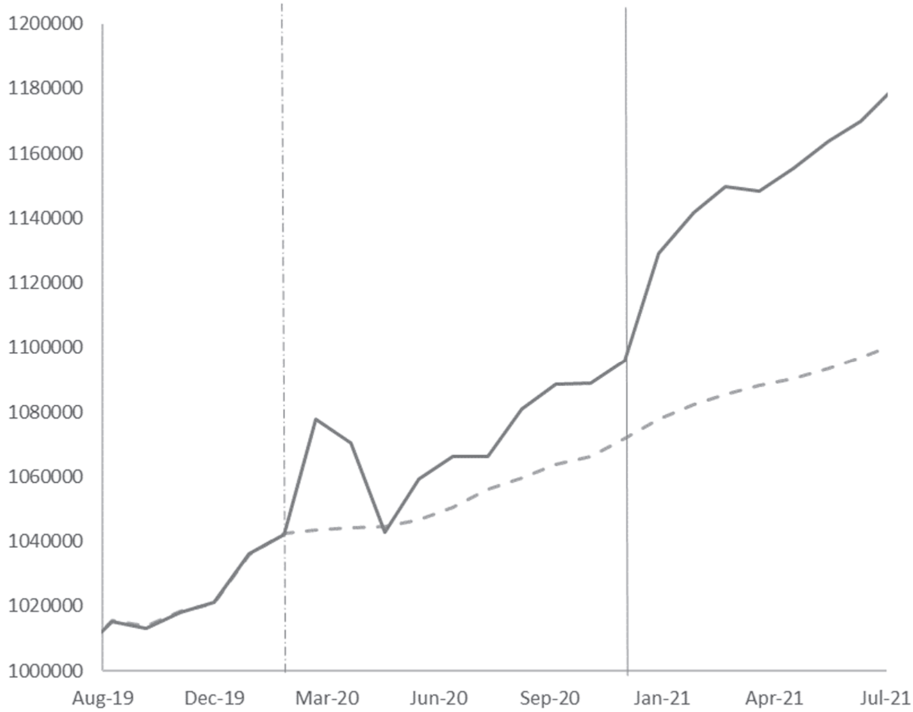
Figure 4. Impulse response Function Following the COVID-19 Pandemic for Retail Sales in Albania



5. DISCUSSION

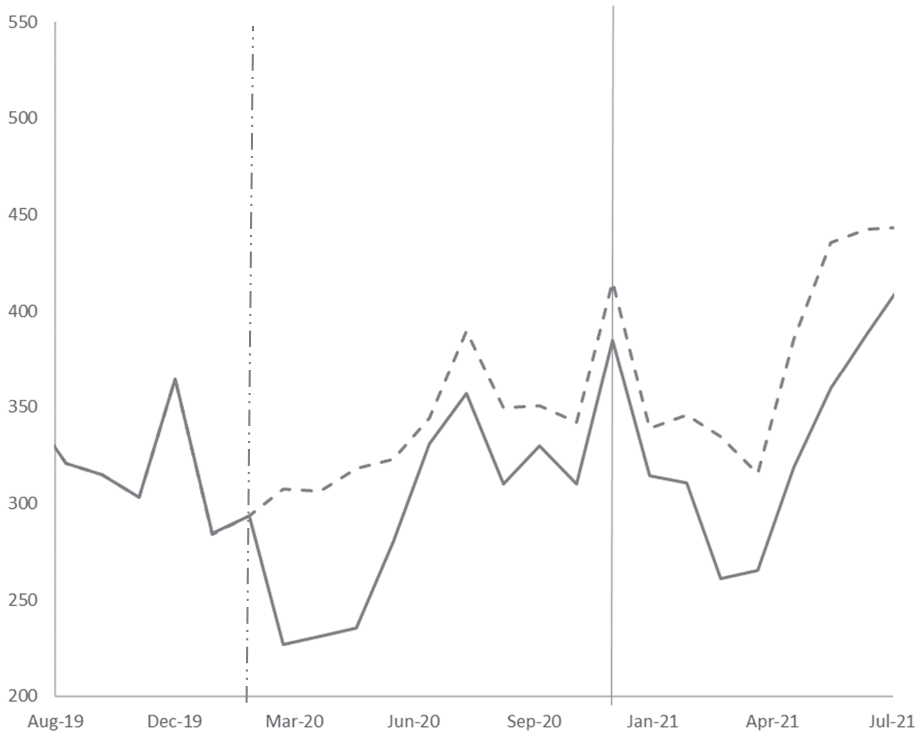
To provide a context for the results reported in the previous section we generated and compared two different sets of forecasts for retails sales and consumer savings. First, we generated forecasts for retails sales and consumer savings using the respective models discussed in the results section using all available data. Second, we generated forecasts using the models estimated with data prior to the start of the pandemic. These forecasts are presented in Figures 5 and 6 for savings and retail sales respectively. Figure 5 shows that as a result of the pandemic, consumer savings are expected to increase much more than expected prior to the pandemic. Figure 6 shows that while retail sales in Albania are expected to bounce back and increase, their rate of increase is still below the expected rate of increase prior to the pandemic.

Figure 5. Forecasts of Savings (Billion Albanian Lek).



Note: The solid line uses all available data, while the dashed line uses data up to February 2020 – prior to the start of the pandemic. The vertical lines indicate where the out-of-sample forecasts start, for each series, respectively.

Figure 6. Forecasts of Retail Sales Index (March, 2005=100).



Note: The solid line uses all available data, while the dashed line uses data up to February 2020 – prior to the start of the pandemic. The vertical lines indicate where the out-of-sample forecasts start, for each series, respectively.

These results provide evidence for both policymakers and the Bank of Albania with regard to fiscal and monetary policy response. It appears that to overcome pandemic effects, accommodating fiscal and/or monetary policies may be needed to support an increase in consumer spending. However, such policies need to be well targeted and probably temporary, especially for countries like Albania that face financing/borrowing constraints.

Governments across the globe were fast in acting to protect their people and businesses amidst what is likely to be the biggest recession of our time. The strategy and scale of economic interventions (e.g., fiscal response) have been broad, ranging between 2.5% and 50% of GDP (Danielli et al. 2021). The scale of intervention was related to the impact of COVID-19 on the effected economy (for example, in Italy the percentage of GDP-equivalent fiscal response was one of the highest, at 50% (IMF 2021)) and available resources.

In Albania the government response has been at the low end of the scale: 2.8% of GDP in 2020 in budget spending, sovereign guarantees, and tax deferrals, and 1% in 2021, mainly in budget spending for wage increases for medical staff, unemployment benefits, and social assistance (IMF 2021). Based on the results presented here, this scale of intervention appears insufficient and a stronger response is needed from the Albanian government to mitigate the negative economic effects of the pandemic.

6. CONCLUSION

This paper investigates the effects of the COVID-19 pandemic on savings and consumer behaviour in Albania, a country which is both among those most affected by COVID-19 and one of the poorest countries in Europe. Albania has experienced a traumatic and prolonged transition from a planned to a free market economy. This transition has been characterized by high levels of social, economic, and political volatility. As an example, previous research (Lami et al. 2014) has found that Albanian households' consumption spending decreases before elections because of increased uncertainty about the future economic situation as a result of the election. The perceived risk of unemployment, which looms larger in countries with high levels of unemployment such as Albania, is found to have a strong direct impact on current household spending and saving behaviour (ibid). Therefore, Albania is an interesting case with which to assess the impact of COVID-19 on households' consumption and saving behaviours.

Intervention analysis (Enders 1995) was employed to measure the effects of the pandemic on two key macroeconomic measures for the Albanian economy, consumer savings and retail sales. The results of this study show that the COVID-19 pandemic caused both short-run and long-run increases in savings among Albanian citizens. Additionally, the findings show that in Albania the COVID-19 pandemic caused both a short-run decrease and a long-run increase in retail sales. Lower consumption spending is also associated with the cutback in sources of livelihood and the aftermath of reduced domestic consumption: malnutrition is one of the most concerning issues for Albania, considering the highly vulnerable population and the already existing high poverty rate. Nevertheless, the Albanian government response has been relatively weak – a stronger response is needed to mitigate the negative economic effects of the pandemic.

One limitation of this study is that it focuses on a single country. However, the findings can be considered relevant to other emerging transition or developing economies. An additional limitation is the limited time span of the pandemic effect due to the time lag between the effect and the availability of official statistics.

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DOES FOREIGN DIRECT INVESTMENT SPUR ECONOMIC GROWTH? NEW EMPIRICAL EVIDENCE FROM SUB-SAHARAN AFRICAN COUNTRIES

ABSTRACT: *In this study we re-examine the relationship between foreign direct investment (FDI) and economic growth in 27 sub-Saharan African (SSA) countries during the period 1990–2019. Unlike some previous studies, we clustered SSA countries into two groups, namely low-income and middle-income countries. We also employed three panel data techniques in a stepwise fashion, namely the dynamic ordinary least squares (DOLS), the fully modified ordinary least squares (FMOLS), and heterogeneous Granger non-causality approaches. Our results show that while the positive impact of FDI on economic growth is supported by both DOLS and FMOLS techniques in low-income countries, in middle-income countries only the DOLS technique supports this finding. This shows that the impact of FDI may be sensitive to the level of income of the recipient country. Overall, the results show that FDI inflows play a larger role in stimulating economic growth in low-income SSA countries than in middle-income SSA countries. These findings are also corroborated by heterogeneous Granger non-causality results. However, these findings are not surprising, given that many low-income countries tend to be more dependent on inward FDI inflows to stimulate their economic growth than middle-income countries. Policy recommendations are discussed.*

KEY WORDS: *FDI, economic growth, sub-Saharan African countries, panel data analysis*

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1. INTRODUCTION

The relationship between FDI and economic growth has been investigated extensively in the literature. Both the neoclassical and endogenous growth models have shown that there is a symbiotic relationship between FDI and economic growth. Findlay (1978), for example, argues that FDI can lead to an increase in the rate of technological progress in the host country in a number of ways, namely management practices and a 'contagion' effect, which arises owing to the adoption of more-advanced technology from the source country (see also Odhiambo, 2021). In other studies it has been found that FDI increases the accumulation of capital in the host country by providing it with new technologies and inputs (see Blomström, Lipsey & Zejan 1996; Borenztein, DeGregorio & Lee, 1998, amongst others). In addition, FDI may serve as a source of productivity gains to the domestic firms of the host country through the spillover effect (see Chanegriha et al., 2020). As an example, studies have shown that multinational corporations (MNCs) impact positively on human capital through the training of unskilled and skilled labour (see Anwar & Nguyen, 2010). Research and development activities, which are usually undertaken by MNCs, may also result in the growth of human capital in host countries, which may eventually boost their economic growth in the long run (see Blomström & Kokko 2001). Technology transfers from MNCs to host countries have been found to be the most important channels through which the presence of MNCs creates positive externalities in the host countries (see OECD, 2002). FDI has also been linked not only to more efficient productive methods, but also to efficient management (see Escobari & Vacaflores, 2015). FDI can, inter alia, lead to direct or indirect job creation, an increase in exports, and an improvement in levels of technology, thereby leading to economic growth (Jordaan, 2012). In some studies it has also been argued that FDI is an important channel by which technology can be transferred to the recipient country, since it has been found in some countries that the FDI contribution to growth is higher than the domestic investment (see Borensztein et al., 1998). In summary, theoretical arguments on the link between FDI and growth can be broadly discussed from four viewpoints. These are the Modernisation Theory (see Calvo & Sanchez-Robles, 2002; Kumar & Pradhan, 2002; and Nath, 2005), the Dependency Theory (see Bornschier & Chase-Dunn, 1985; and Amin, 1974), the Neoclassical Growth Theory (Solow, 1956; and Swan, 1956), and the Endogenous Growth Theory (Romer, 1986; and Lucas, 1988). According to the endogenous theories, based on Romer (1986) and Lucas (1988), FDI involves the transfer of technology as well as the training of labour, which contribute significantly to human capital accumulation, thereby inducing technological progress and long-term growth (see also Tsaurai & Odhiambo, 2012).

Although a plethora of studies exist on the FDI–growth nexus in some developing countries, most of the previous studies have focused mainly on Asian countries (see, for example, Baharumshah & Almasaied, 2009; Wang, 2009; Hoang et al.,2010; Muhammad & Khan, 2019; and Ang, 2009, amongst others). In addition, most previous studies focus mainly on either the causal relationship between FDI and growth or on the impact of FDI on growth. Very few studies have gone the full distance to examine both the impact and the causal relationship between FDI and economic growth in SSA countries. Moreover, some of the previous studies suffer from methodological weakness. As an example, some studies use cross-sectional data, which may not fully explore the dynamic relationship between FDI and economic growth. The weaknesses of cross-sectional data have been extensively discussed in the literature (see Odhiambo, 2008; Ghirmay, 2004; Quah, 1993; Casselli et al., 1996). By lumping together countries which are at different levels of development, the cross-sectional approach fails to account for country-specific effects inherent in the FDI–growth nexus. Moreover, it has been shown that the cross-section approach may produce inconsistent and misleading estimates owing to the potential biases that may arise owing to the existence of heterogeneity among the study countries (see also Odhiambo, 2008; Ghirmay, 2004). Even in instances where a panel dataset has been used, some of the methodological weaknesses associated with a long panel dataset, such as cross-sectional dependency, have not been addressed fully. Other studies also over-rely either on fixed effects or random effects panel estimation techniques, which may not account for the endogeneity inherent in panel data.

The current study is therefore aimed at innovatively addressing some of these weaknesses by examining the FDI–growth nexus in 27 SSA countries[1] using a wide range of estimation techniques. In addition, the study divides the studied

[1] The countries used in this study are Sierra Leone, Burkina Faso, Burundi, Sudan, Chad, Madagascar, Malawi, Mali, Togo, Mozambique, Gambia, Niger, Rwanda, Kenya, Benin, Senegal, Tanzania, Cameroon, Comoros, Republic of the Congo, Cote d'Ivoire, Eswatini, Nigeria, Ghana, Gabon, Botswana, South Africa.

countries into two groups, namely a low-income group comprising 13 countries and a middle-income group comprising 14 countries. The selection of the study countries and the study period were largely driven by the availability of data for low-income and middle-income SSA countries. This study is timely, given the proposed African Continental Free Trade Area (AfCFTA). [2] The AfCFTA will involve deep reforms necessary to enhance long-term growth in African countries, which include cutting the red tape and simplifying customs procedures. The implementation of AfCFTA is therefore expected to boost economic growth in Africa, reduce poverty, and broaden its economic inclusion. According to the current projections, the implementation of AfCFTA is expected to boost Africa's income by $450 billion by 2035 (World Bank, 2020).

To address the limitations of previous studies, the dynamic ordinary least squares (DOLS), the fully modified ordinary least squares (FMOLS), and the heterogeneous non-causality approaches have been used in a step-wise fashion. Other tests, such as cross-section dependence tests, have also been incorporated into the current study. Four tests are used to test for cross-section dependence, namely Breusch-Pagan LM, Pesaran scaled LM, Baltagi et. al. (2012), and Pesaran CD. Given the weaknesses associated with first-generation unit root tests in the presence of cross-section dependence, the current study has used the second-generation unit root tests associated with Bai and Ng (2004), Panel Analysis of Non-stationarity in Idiosyncratic and Common components (PANIC), and Pesaran (2007) CIPS (cross-sectionally augmented IPS) alongside the first-generation unit root tests to examine whether the variables used in this study are conclusively I(0) or I(1).

The study closest to this study is Opoku et al. (2019). However, the current study differs fundamentally from Opoku et al. (2019) in several ways. First, while Opoku et al. (2019) focus mainly on the sectoral transmission channels by which FDI affects growth, the current study focuses on the intertemporal relationship between FDI and growth using both the impact model and the causality model. Second, unlike Opoku et al. (2019) who mainly used system GMM, the current study uses DOLS, FMOLs, and heterogenous Granger-causality to examine the nexus between FDI and growth, while accounting for cross-sectional dependence using the second-generation unit root and cointegration tests. Third, unlike

[2] See also Ofari & Asongu (2022).

Opoku et al. (2019), the current study divides SSA data into two groups, namely low-income and middle-income groups, to examine whether the relationship between FDI and economic growth depends on the countries' income level. To our knowledge, this may be the first study of its kind to empirically examine, in detail, the nexus between FDI and growth in SSA countries using disaggregated data and an array of modern panel data techniques.

The remainder of the paper is organised as follows. Section 2, reviews the literature on the impact of FDI on economic growth. Section 3 presents the methodology and the empirical model specification in a step-wise fashion. Section 4 deals with the empirical analysis, as well as a discussion of the results. The study concludes in section 5.

2. EMPIRICAL LITERATURE

Studies that have been conducted on the relationship between FDI and growth can be broadly clustered into two groups, namely studies whose findings are consistent with a positive relationship between FDI and growth, and studies in which the findings support a mixed, negative, or insignificant relationship. Adams (2009), for example, while examining the link between FDI, domestic investment, and economic growth in SSA using OLS and fixed-effects estimation techniques during the period 1990–2003, finds that FDI correlates positively with economic growth in the OLS model, but only after controlling for country-specific effects. Baharumshah & Almasaied (2009) examine the role of FDI in economic growth in Malaysia. Using data from 1974 to 2004, the study finds FDI to have a positive impact on growth; however, its impact is found to be smaller than that of non-FDI investment. While analysing the relationship between FDI and economic growth in 79 countries during the period 1980–2003, Batten & Vo (2009) find that the impact of gross FDI flow on economic growth is stronger in countries that have a higher level of educational attainment. The same results apply to countries that are more open to international trade and have higher levels of stock market development. Wang (2009), using data from 12 Asian economies over the period 1987–1997, finds strong evidence showing that manufacturing sector FDI has a positive impact on growth in the host countries. Hoang et al. (2010), using the panel data model to examine the impact of FDI on growth in Vietnam's 61 provinces during the period 1995–2006, find FDI to have a strong positive impact on growth. Nistor (2014), using Romanian data from 1990 to

2012, finds that FDI inflows have a positive impact on gross domestic product in the country under study. Adams & Opoku (2015), while examining the effect of FDI on economic growth and how regulatory regimes affect the FDI–growth relationship in SSA using the GMM estimation technique, find that neither FDI nor regulations have a significant effect, but their interaction has a significant positive effect on economic growth. Pegkas (2015), in examining the link between FDI and growth in Eurozone countries in 2002–2012, finds that, consistent with some theoretical explanations, FDI is a significant determinant of economic growth. While using various panel data estimation techniques, Muhammad & Khan (2019) find, inter alia, that FDI inflows to Asian host countries have a positive impact on growth. Nketiah-Amponsah & Sarpong (2019), investigating the impact of infrastructure and FDI on growth using data from 46 SSA countries, find that FDI enhances economic growth only when interacting with infrastructure. Opoku et al. (2019) examine the relationship between FDI, sectoral effects, and economic growth in 38 African countries during the period 1960–2014. Using a system GMM, the study finds that although FDI has an unconditional positive impact on economic growth, its growth-enhancing impact becomes imaginary with the introduction of conditional sectoral effects. Pradhan et al. (2017), in investigating the causal relationship between trade openness, foreign direct investment, financial development, and economic growth in 19 Eurozone countries during the period 1988–2013, find that FDI inflows have propelled economic growth in the studied countries in the short run. Pradhan et al. (2018) examine the interactions between the diffusion of mobile phones, foreign direct investment, financial development, ICT goods imports, and economic growth in the G-20 countries during the period 1990–2014. Using a multivariate framework, the study finds, inter alia, that there is a long-run unidirectional causality from foreign direct investment to economic growth in the studied countries. Pradhan et al. (2019), examining the heterogeneous relationship between financial development, foreign direct investment (FDI), and economic growth using a sample of G-20 countries over the period 1970–2016, find that both FDI and financial development matter in the determination of long-run economic growth in the studied countries. Asongu & Odhiambo (2020), examining the relationship between FDI, ICT, and economic growth in 25 sub-Saharan African countries using the GMM approach, find that both internet penetration and mobile phone penetration overwhelmingly modulate FDI to induce overall positive net effects on all three economic growth dynamics.

Ibhagudi (2020), examining the effect of FDI on economic growth in sub-Saharan African countries using a threshold regression framework, finds that FDI accelerates economic growth when SSA countries have achieved certain threshold levels of inflation, population growth, and financial market development. More recently, Arvin et al. (2021), examining the links between ICT connectivity and penetration, trade openness, foreign direct investment, and economic growth using data from the G-20 countries during the period 1961–2019, find inter alia that economic growth is dependent on FDI in the long run in the studied countries.

Apart from the above-mentioned studies, a few studies cast doubt on the positive impact of FDI on economic growth. These studies find the relationship between FDI and economic growth to be mixed, negative, or not significant at all. These include studies such as Eller et al. (2006), Ang (2009), Alvarado et al. (2017), and Carbonell & Werner (2018), among others. Eller et al. (2006), for example, while analysing the effect of financial sector FDI on growth through the efficiency channel using data from 11 Central and Eastern European countries, find FSFDI to have a hump-shaped impact on economic growth in the studied countries. Ang (2009) examines the roles of FDI and financial development in economic development in Thailand during the period 1970–2004. Using the unrestricted ECM estimator, the study finds that while financial development stimulates economic development, FDI negatively impacts output expansion in the long run. Alvarado et al. (2017) examine the impact of FDI on growth in 19 countries in Latin America during the period 1980–2014. Using panel data econometric techniques, the study fails to find any clear direction in the impact of FDI on growth in the studied countries. Moreover, the study finds that the impact of FDI on economic growth is sensitive to the countries' level of development. Carbonell & Werner (2018), using data from Spain during the 1984(Q1)–2010(Q4) period, fail to find any evidence which shows that FDI stimulates growth. The authors attribute this finding to the fact that the bulk of Spanish FDI inflows are from foreign takeovers which are largely in the construction sector.

3. METHODOLOGY

In this study, panel data have been used to analyse the nexus between FDI and growth in SSA. The advantages of using panel data have been covered extensively in the literature (see Rahman et al., 2021).

The panel model employed to analyse the relationship between FDI and economic growth in the selected SSA countries can be expressed as follows:

$$\mathrm{Y_{it}} = \gamma_{\mathrm{it}} + \delta_{it} + \beta_{1i} FDI_{it} + \beta_{2i} Trade_{it} + \beta_{3i} Labour_{it} + \beta_{4i} GFCF_{it} + \mu_{it} \tag{1}$$

where *i* refers to cross-sectional observation, *t* indicates the time period, *Y* = per capita GDP, *FDI* = Foreign direct investment (% of GDP), *Trade* = Exports + Imports (% of GDP), *Labour* = Labour force, *GFCF* = Gross fixed capital formation, $\delta_{it} and \beta_{1i}$ = country-specific effects and deterministic trend effects, respectively, and μ_{it} = error term.

In this paper we use real GDP per capita as a proxy for economic growth. This proxy has been used extensively in the literature (see Asongu, 2013; Odhiambo, 2014, 2022; Asongu et al., 2022;). FDI, on the other hand, is measured by FDI inflows as a percentage of GDP (see Asongu et al., 2020; Odhiambo, 2021). According to the attendant literature, FDI is expected to spur economic growth inter alia through technology diffusion and increases capital accumulation in the host country through the introduction of new inputs and technologies (see Blomstrom et al., 1992; Borenszteinet al., 1998; Odhiambo, 2021). Consequently, the coefficient of the FDI is expected to be positive and statistically significant. The control variables used in this study are trade, labour, and gross fixed capital formation. The justification for including these variables was informed by both the theoretical and empirical literature. The inclusion of trade in the growth equation is informed by the role that trade plays in economic growth and development. An increase in trade is expected to have a positive impact on economic growth. Put slightly differently, trade has been found to be an engine of development (see Frank, 1968). Hence, the coefficient of trade is expected to be positive and statistically significant. The impact of labour productivity on economic growth has also been supported by the attendant literature. In particular, labour quality has been found to have a positive impact on economic growth, as countries with higher labour quality are likely to be associated with higher productivity growth (see Barro, 2001). Consistent with extant literature, an increase in gross fixed capital formation is expected to lead to an increase in economic growth as it leads to more jobs and hence an increase in employment. Consequently, the coefficient of gross fixed capital formation is expected to be positive and statistically significant (see also Levine & Renelt, 1992; Mankiw et al., 1992).

Consistent with previous studies, the model presented in Equation (1) can be estimated using the DOLS. The advantage of the DOLS is that it can correct endogeneity, serial correlation, and simultaneity problems via differenced leads and lags (see also Maji et al., 2019). In this way, the DOLS can generate an unbiased estimate (see Mc-Coskey & Kao, 1998; Kao & Chiang, 2000; Maji et al., 2019). For robustness check, the FMOLS has also been applied alongside the DOLS in this analysis. The main difference between DOLS and FMOLS relates to how the autocorrelation is corrected in the regression. FMOLS, for example, is regarded as a nonparametric correction that adjusts for autocorrelation (Bellocchi et al., 2021). DOLS, on the other hand, which has been found to outperform both the OLS and FMOLS estimators, allows for the addition of more lagged and lead variables in the regression. A summary of the variables used in this study is presented in Table 1.

Table 1: Variable Description, Expectations, and Sources

Variable	Description	Expectation	Source
y/N	Economic growth	NA	WDI
FDI	Foreign direct investment	+	WDI
Trade	Total trade	+	WDI
LABOUR	Labour force	+	WDI
GFCF	Gross fixed capital formation	+	WDI

Heterogenous Granger Causality

The heterogeneous panel Granger non-causality estimator, based on Dumitrescu & Hurlin (2012), is used to examine the causal relationship between FDI and economic growth. The advantage of this technique is that it considers the CSD ratio. It has also been found to account for both the time dimension and the size of cross-sections. The Dumitrescu & Hurlin (D-H) panel Granger non-causality model can be expressed as follows:

$$y_{it} = \alpha_i + \sum_{k=1}^{K} \delta_i^k y_{i(t-k)} + \sum_{k=1}^{K} \beta_i^k x_{i(t-k)} + \varepsilon_{i,t} \tag{2}$$

where *y* and *x* = variables, t = time dimension, i.e., *t* = 1……………..T, and *i*= individual, i.e., *i* = 1…………………………N.

Based on D-H, the null hypothesis of no causality for each panel group (H_0: $\beta_i = 0, i = 1,2 \ldots\ldots\ldots\ldots, N)$) is tested against the alternative hypothesis of causality between the variables within the panel group for each country (i.e., H_1: $\beta_i = 0, i = 1,2, \ldots\ldots\ldots\ldots, N; \beta_i \neq 0;\ i = N_1 + 1, N_1 + 2, \ldots\ldots\ldots\ldots, N$).

The study employs annual data from 1990–2019. The data were sourced from the World Bank's World Development Indicators. The World Bank data were supplemented by national databases.

4. EMPIRICAL ANALYSIS

4.1 Cross-Section Dependence

Before proceeding with the unit root test, it is important to conduct a panel cross-section dependence test in order to account for possible cross-section dependence among the countries under study. Cross-section dependence could result from factors such as international trade, financial integration, and globalisation, which may result in external shocks from other countries (see Chang et al., 2013). Studies have also shown that ignoring cross-section dependency in a panel estimation can have serious consequences as it may lead to substantial bias and size distortions (Pesaran, 2006). For this reason, four tests for cross-section dependence have been employed to test the existence of cross-section dependence in the estimation. These are Breusch-Pagan LM, Pesaran scaled LM, Bias-corrected scaled LM, and Pesaran CD. The results of the cross-section dependence test are reported in Table 1.

Table 1: Cross-section dependency tests

	Cross-section dependency results			
Series	Breusch-Pagan LM	Pesaran scaled LM	Bias-corrected scaled LM	Pesaran CD
	Low-income countries (LICs)			
y/N	988.2482*** (0.0000)	72.8782*** (0.0000)	72.6541*** (0.0000)	15.7539*** (0.0000)
FDI	369.0149*** (0.0000)	23.2998*** (0.0000)	23.0757*** (0.0000)	13.1697*** (0.0000)
Trade	456.8627*** (0.0000)	30.3333*** (0.0000)	30.1092*** (0.0000)	13.3799*** (0.0000)
GFCF	346.0813*** (0.0000)	21.4637*** (0.0000)	21.2396*** (0.0000)	9.31481*** (0.0000)
Labour	1078.9830*** (0.0000)	80.1428*** (0.0000)	79.9186*** (0.0000)	7.7265*** (0.0000)
	Middle-income countries (MICs)			
y/N	1490.4070*** (0.0000)	103.7309*** (0.0000)	103.4896*** (0.0000)	24.6420*** (0.0000)
FDI	450.3611*** (0.0000)	26.6376*** (0.0000)	26.3962*** (0.0000)	15.1154*** (0.0000)
Trade	517.8897*** (0.0000)	31.6432*** (0.0000)	31.4018*** (0.0000)	14.3418*** (0.0000)
GFCF	395.0901*** (0.0000)	22.5407*** (0.0000)	22.2993*** (0.0000)	10.2377*** (0.0000)
Labour	1197.9120*** (0.0000)	82.0498*** (0.0000)	81.8084*** (0.0000)	8.2117*** (0.0000)

The results reported in Table 1 for LICs and MICs show that the four cross-section dependence tests have largely rejected the null hypothesis of no cross-section dependence in both LIC and MIC. This indicates that there is cross-section dependence in the data used. These results suggest the use of second-generation unit root tests in order to account for the presence of cross-section dependency.

4.3 First- and second-generation panel unit root tests

Having detected the presence of cross-section dependence, it is important to use the second-generation panel unit root tests together with the first-generation tests when conducting unit root tests. The results of the stationarity tests are reported in Tables 2 and 3, respectively.

Table 2: The results of the first-generation panel unit root tests

	Low-Income SSA Countries				**Middle-Income SSA Countries**			
	LLC t-Statistics		**IPS**		**LLC t-Statistics**		**IPS**	
	Level	**First Difference**	**Level**	**First Difference**	**Level**	**First Difference**	**Level**	**First Difference**
y/N	1.5353	–12.1949***	–0.5198	–2.9640***	–1.0270	–5.4511***	–0.5198	**–2.9649***
FDI	–0.5571	–9.0254***	0.3120	–3.6716***	–0.5571	–9.0254***	0.3120	**–3.6716***
Trade	–0.7709	–6.0932***	–0.9676	–5.2117***	–0.7709	–6.0932***	–0.9676	**–5.2117***
Labour	0.4763	–6.6544***	–0.1662	–8.1589***	–0.5341	–5.2148***	1.1085	**–3.5198***
GFCF	–0.7540	–8.4388***	–1.0569	–10.1556***	–1.0768	–6.7191***	–1.1018	**–10.5788***

Note: *** indicates rejection of the respective null hypothesis at the 1% significance levels, respectively.

Table 3: The results of second-generation panel unit root tests

	Low-Income SSA Countries				**Middle-Income SSA Countries**			
	Bai and Ng – PANIC		**Pesaran – CIPS**		**Bai and Ng – PANIC**		**Pesaran – CIPS**	
	Level	**First Difference**	**Level**	**First Difference**	**Level**	**First Difference**	**Level**	**First Difference**
y/N	–0.5588	–2.1440**	–0.7176	–3.8776***	0.5445	–2.8226***	–1.4741	**–2.8796**
FDI	1.3460	–2.0235**	1.2430	–6.1520***	1.5756	–2.6100***	–1.6391	**–3.3036***
Trade	1.1204	–2.0120**	–1.4500	–6.6349***	–1.3304	–2.2770**	–0.2481	**–3.2435***
Labour	0.3172	–3.4070***	–1.5941	–3.1816***	–0.8297	–2.0599**	–0.7854	**–3.6498***
GFCF	–0.3135	–2.8151***	–1.0217	–5.1948***	0.4673	–2.5046**	–0.7706	**–3.7075***

Note: ** and*** indicate rejection of the respective null hypothesis at the 5% and 1% significance levels, respectively.

The results of the first generation unit root tests reported in Table 2 show that the variables are conclusively I(1). These results have also been confirmed by the second-generation unit root tests reported in Table 3, which show that both

PANIC and Pesaran–CIPS reject the stationarity in levels in favour of stationarity in the first difference.

4.4 Panel Cointegration Test

Since the variables included in this study have been found to be I(1), it is important to test whether the variables y/N, FDI, Trade, Labour and Inv are cointegrated. For this purpose, the study uses three tests, namely the Pedroni (1999; 2004), the Kao (1999), and the Westerlund (2005) class of tests. The results of the Pedroni, Kao, and Westerlund cointegration tests are reported in Table 4, panels 1, 2, and 3, respectively.

Table 4: Panel cointegration results

PANEL 1: Pedroni cointegration test				
	Low-income countries		**Middle-income countries**	
	Statistic	**Probability**	**Statistic**	**Probability**
Pedroni panel cointegration test – within-dimension				
Panel v-Statistic	–1.9318	0.9733	–2.3918	0.9916
Panel rho-Statistic	–0.4581	0.3235	–1.6736	0.0471
Panel PP-Statistic	–9.0726	0.0000	–27.1338	0.0000
Panel ADF-Statistic	–4.8301	0.0000	–12.9086	0.0000
Pedroni panel cointegration test – between-dimension				
Group rho-Statistic	1.0818	0.8603	0.3157	0.6239
Group PP-Statistic	–17.9627	0.0000	–19.2619	0.0000
Group ADF-statistic	–6.2293	0.0000	–5.5543	0.0000
PANEL 2: Kao residual cointegration Test				
	Low-income countries		**Middle-income countries**	
	Statistic	p-value	Statistic	p-value
ADF	–4.2862	0.0000	–3.0510	0.0011
PANEL 3: Westerlund (2005) cointegration Test				
	Low-income countries		**Middle-income countries**	
	Statistic	p-value	Statistic	p-value
Variance ratio	–3.4880	0.0002	–3.2378	0.0006

The results of the Pedroni cointegration test reported in Table 4 (Panel 1) show that all the variables included in our model for LICs and MICs are cointegrated. This is confirmed by the Panel PP-Statistic, Panel ADF-Statistic, Group PP-

Statistic, and Group ADF-Statistic, which are all significant at the 1% level in LICs and MICs. In other words, the results show that four of the seven Pedroni residual cointegration tests confirm that the variables are cointegrated in both income groups. The Kao (1999) and Westerlund (2005) tests reported in panels 2 and 3 also show that the variables are cointegrated in both income groups. This finding is confirmed by the ADF statistics in the Kao cointegration test and the variance ratio in the case of the Westerlund (2005) test, which are both found to be statistically significant.

4.5 Dynamic OLS (DOLS) and fully modified OLS (FMOLS)

In this section, DOLS and FMOLS are used to examine the impact of FDI on economic growth in LICs and MICs. These two techniques account for endogeneity and serial correlation. The results of DOLS and FMOLS are reported in Table 5.

Table 5: DOLS and FMOLS results

Explanatory variable	Low-income countries (LICs)				Middle-income countries (MICs)			
	DOLS		FMOLS		DOLS		FMOLS	
	Coefficient	t-statistic	Coefficient	t-statistic	Coefficient	t-statistic	Coefficient	t-statistic
FDI	1.5753***	5.0490	4.5333**	2.1979	4.9207**	2.4436	−4.3580	−1.4007
TRADE	0.4967***	4.3156	1.7594***	3.6309	−3.6249***	−12.4217	−3.0713***	−3.4221
LABOUR	2.3170**	2.1507	5.4579***	17.3822	14.7700***	6.1535	24.2382**	2.3384
GFCF	−0.0626	−0.3673	4.9157***	4.8073	11.5269***	12.6763	8.6798***	4.7962

Note: *** and ** indicate significance at 1% and 5% levels, respectively

The results in Table 5 clearly show that the impact of FDI on economic growth in SSA countries is not unanimous and depends on the countries' income level. In the case of LICs, FDI is found to have a positive impact on growth when DOLS and FMOLS are used as estimators. This finding is confirmed by the coefficient FDI in the economic growth equation, which is found to be positive and statistically significant in the DOLS and FMOLS panels. Specifically, the coefficient of FDI is positive and statistically significant at 1% and 5% in the DOLS and FMOLS panels, respectively. Contrary to the results for the LICs, the results for MICs show that FDI has a positive impact on growth only when the estimations are conducted using DOLS estimators. This finding is evidenced by the coefficient of the FDI, which is positive and statistically significant in the

DOLS panel, but not in the FMOLS panel. This finding shows that while it can be concluded that FDI has an overall positive impact on economic growth in many SSA countries, LICs tend to benefit more from FDI than MICs. While the overall positive impact of FDI on economic growth is contrary to studies such as Ang (2009) for the case of Thailand, it is consistent with studies such as Baharumshah & Almasaied (2009) for the case of Malaysia, and Hoang et al. (2010) for the case of Vietnam, among others. The remaining results show that in the case of LICs, trade and labour have a positive impact on economic growth in DOLS and FMOLS models. This finding is supported by the coefficients of trade and labour in the growth equation, which are positive and significant at the 1% level in the DOLS and FMOLS panels. Unlike in the case of trade and labour, gross fixed capital formation has been found to have a positive impact on economic growth only when FMOLS is used as an estimator, but not when DOLS is used. This has been supported by the coefficient of gross fixed capital formation, which is positive and statistically significant in the DOLS specification, but not in the FMOLS specification.

For MICS, the results indicate that trade has a negative effect on economic growth, irrespective of whether DOLS or FMOLS is used as an estimator. This is supported by the coefficient of trade in the economic growth equation, which is negative and significant in the DOLS and FMOLS panels. This finding, though contrary to our expectation, is consistent with some previous studies, such as Rigobon & Rodrik (2005), who find openness (trade/GDP) to have a negative impact on income. This finding is also unsurprising, considering the nature of the trade balance in some SSA countries. Since some SSA countries are largely operating on a trade deficit rather than a trade surplus, it is likely that the cumulative deficit accrued over time by some middle-income countries could have a negative bearing on their economic growth trajectories, thereby leading to a negative relationship between trade and growth. The results also show that in MICS, labour and gross fixed capital formation have a positive impact on economic growth in the DOLS and FMOLS models. This is confirmed by the coefficients of labour and gross fixed capital formation in the economic growth equation, which are found to be positive and statistically significant in both DOLS and FMOLS panels.

4.6 Heterogeneous Panel Causality Analysis

In this study, the Dumitrescu-Hurlin (D-H) panel Granger-causality model, which supports cross-sectional heterogeneity, is used to explore the causal relationships between FDI and economic growth, as well as other variables included in the economic growth model. Since the D-H panel Granger-causality test requires the data series to be stationary, we have to convert our series into the first difference. The results of the Granger-causality between FDI and economic growth, and other variables are summarised in Table 6.

Table 6: Heterogeneous panel causality test

	Low-income countries (LICs)			Middle-income countries (MICs)		
Null Hypothesis:	Zbar-Stat.	Prob.	Causality	Zbar-Stat.	Prob.	Causality
DFDI does not homogeneously cause Dy/N	2.3752	0.0175	DFDI→ Dy/N	1.0054	0.3147	DFDI [0] Dy/N
Dy/N does not homogeneously cause DFDI	−1.5763	0.1149		−0.4279	0.6687	
DGFCF does not homogeneously cause Dy/N	0.4460	0.6556	DGFCF [0] Dy/N	−0.0390	0.9689	Dy/N →DGFCF
Dy/N does not homogeneously cause DGFCF	−0.5347	0.5928		1.8687	0.0617	
DLABOR does not homogeneously cause Dy/N	0.0234	0.9813	Dy/N→DLABOUR	1.9303	0.0536	Dy/N ↔ DLABOUR
Dy/N does not homogeneously cause DLABOUR	1.8218	0.0685		2.0668	0.0388	
DTRADE does not homogeneously cause Dy/N	1.1349	0.2564	DTRADE [0] Dy/N	0.4546	0.6494	Dy/N→DTRADE
Dy/N does not homogeneously cause DTRADE	−0.1659	0.8682		1.9765	0.0481	

The empirical results reported in Table 6 show that the causality between FDI and growth in SSA is sensitive to the income group of the studied countries. For LICs there is a distinct unidirectional causality from FDI to economic growth. This finding is confirmed by the Zbar-Statistic, which is significant in the economic growth equation but not in the FDI equation. For the MICs, the empirical results indicate no causality between FDI and economic growth in either direction. This is confirmed by the Zbar-Statistic, which is insignificant in both FDI and economic growth equations. Other results show that 1) economic growth Granger-causes gross fixed capital formation in MICs, but in LICs there is no causality between the two variables; 2) economic growth Granger-causes labour force participation in LICs, but in MICs there is a bi-directional causality between the two variables; and 3) economic growth Granger-causes trade in MICs, but in LICs there is no causality between the two variables.

5. CONCLUSION

This study examines the relationship between FDI and economic growth in 27 SSA countries during the period 1990–2019. SSA data is divided into two income groups, a low-income group and a middle-income group. To address the weaknesses of some of the previous studies, the DOLS, the FMOLS and the heterogeneous non-causality approaches are used in a stepwise fashion in the study. In addition, cross-sectional dependence is tested using four tests: Breusch-Pagan LM, Pesaran scaled LM, Bias-corrected scaled LM, and Pesaran CD. The results of the study clearly show that the impact of FDI on economic growth differs significantly in LICs and MICs. For the LICs, the results show that FDI has a distinct positive impact on economic growth, irrespective of whether DOLS or FMOLS are used as estimators. However, for the MICs the results are not unanimous. The results show that FDI has a positive impact on economic growth only when the estimation is conducted using the DOLS estimator. The heterogeneous non-causality test based on Dumitrescu & Hurlin (2012) corroborates these results: it shows that while FDI Granger-causes growth in LICs, in MICs there is no causal relationship between FDI and growth. These findings show that LICs benefit more from FDI than MICs, and thus the impact of FDI may be sensitive to the level of income of the recipient country. In the main, the results show that FDI inflows play a larger role in stimulating economic growth in low-income SSA countries than in middle-income SSA countries. This finding is unsurprising, given that many low-income countries tend to be more dependent on inward FDI inflows to stimulate their economic growth than middle-income countries. It is therefore recommended that low-income SSA countries should continue to intensify their investment promotion strategies in order to attract more pro-growth investment, while middle-income countries should devise appropriate polices aimed at ensuring that their FDI inflows are pro-growth and do not substitute their domestic investment.

Although all efforts have been made to make this study analytically defensible, like many other empirical studies, it has some limitations. The main limitation is a lack of adequate and reliable data, which forced the study to restrict its study period as well as the number of countries included. These restrictions played a major role in determining the most appropriate estimation techniques used. It is therefore recommended that future studies consider expanding the horizon of the current study by applying other models, such as non-linear ARDL, to determine how negative and positive FDI shocks affect the dynamic relationship between FDI and economic growth in the studied countries.

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INFLATION AND FINANCIAL DEVELOPMENT IN SUB-SAHARAN AFRICA

ABSTRACT: *This paper analyses the relationship between inflation and financial development in sub-Saharan Africa. This issue is important because the level beyond which the inflation rate affects financial development in sub-Saharan Africa is not known, despite the underdeveloped financial system in the region. This paper presents a model in which inflation is endogenously determined, and uses a dynamic panel threshold approach which takes this factor into consideration to ensure improved results. The pure cross-section method and non-dynamic panel threshold approaches were also utilised. Strong evidence of a negative impact of inflation on financial development is obtained, which increases with a rise in inflation. An average inflation threshold of about 5% is established, below which a positive impact on financial development is found. There is, however, a negative relationship beyond the observed threshold. The paper provides policy recommendations as to how authorities should contain inflation to facilitate financial sector development in the region.*

KEY WORDS: *inflation, financial development, panel threshold techniques, sub-Saharan Africa*

JEL CLASSIFICATION: E31, E4, E5, P44

Declarations

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1. INTRODUCTION

Most African countries are highly dependent on primary products and mineral resources, which makes them different from other countries due to the phenomenon of Dutch disease – where by other sectors like manufacturing are negatively affected mostly because of the appreciation of the local currency which makes exports less competitive (Mulwa, 2017), among other challenges. The poor institutional quality in the region has also been identified as a distinct factor that causes variation in the size and nature of macroeconomic variables. This paper investigates how to ensure sustainable development of the financial sector while paying special attention to inflation in sub-Saharan Africa (SSA). This is particularly needed because, compared to the rest of the world, the banking system provides less credit to firms (IMF, 2016; AfDB, 2013). This paper therefore determines the acceptable inflation level that would enable the development of the ailing SSA financial sector and ensure sustainable economic development.

Financial development has been categorised into financial institutions and financial markets (Sahay et al., 2015). The categories are further sub-divided into depth, access, and efficiency. Due to data limitation for countries in SSA this study only deals with the depth element of the financial institution category. The depth category measures the size and liquidity of the market and encompasses private credit, pension fund assets, mutual funds, and life and non-life insurance premiums, all as a share of GDP. These indicators of financial depth are a good proxy for financial development because the financial stability of most economies strongly depends on them (Sahay et al., 2015).

Paramount in the literature is notable evidence of the expediting effect of financial development on economic growth through credit creation, which allocates resources to the most productive sectors of the economy (Levine et al., 2000; King & Levine, 1993; Schumpeter, 1982; Modigliani & Miller, 1963). Sahay et al. (2015) also acknowledge that financial development promotes financial stability, which reduces the impact of shocks. As such, the development of the financial sector has a direct impact on economic growth. On the other hand, a direct and negative relationship between inflation, financial development, and growth has been linked to the unfavourable economic environment created by high inflation, which reduces investor confidence and leads to reduced financial savings and lower rates of credit creation, and hence to stagnant or declining economic growth (Ndoricimpa, 2017; Kremer et al., 2013; Leshoro, 2012). Inflation also affects growth through its influence on financial development (Ehigiamusoe et al., 2018; Huang et al., 2010; Rousseau & Wachtel, 2002).

The mechanism through which inflation affects financial development is linked to credit market friction in the financial markets. A rise in inflation drives down the real rate of return from the financial markets, leading to credit rationing. This leads to fewer loan disbursements, ineffective resource allocation, and reduced capital investment (Huybens & Smith, 1999). Long-run economic performance and financial sector development can both be negatively affected by the decline in capital formation. On the other hand, lower inflation rates can positively drive development in both the real and financial sectors. This can be supported by the existence of a short-run Philips curve, implying that productivity is associated with a rise in prices emanating from the demand side (demand-pull inflation). This type of inflation is closely linked to the demand for factors of production and consumption demand, which push up prices. During a boom, particularly in the less industrialised SSA economies, brisk business in the economy can also spur the demand for financial services, pushing up their prices as well. Thus, there could simultaneously be a rise in general commodity prices, in increased productivity, and in financial sector development, implying a positive relationship between inflation and financial development.

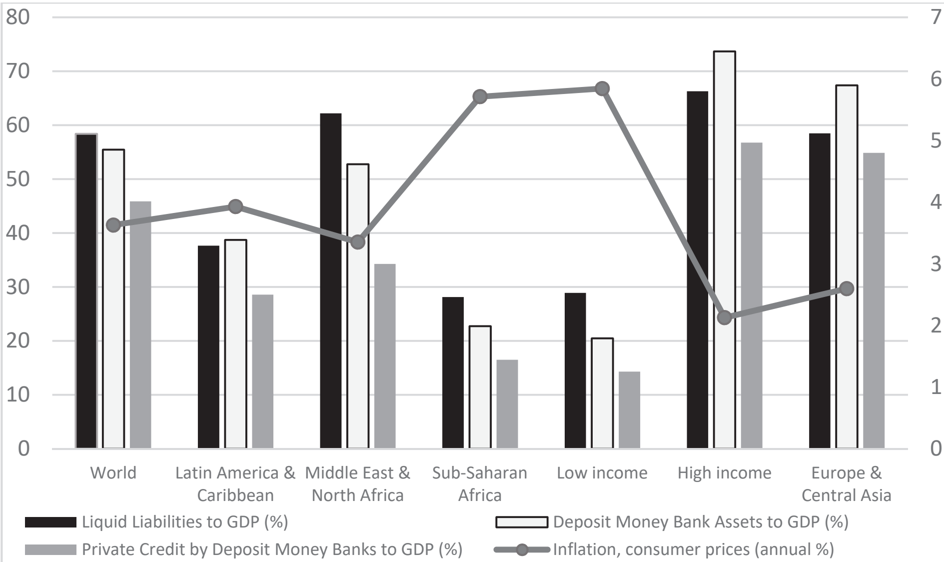
Taking heed of how critical financial sector development could be for development purposes, its connection with inflation in SSA is worth examining. From this perspective, this paper aims to determine the average level of the inflation threshold for financial development in SSA. In contrast to a parallel study by Bandura (2022), which uses the same dataset to determine the indirect impact of inflation on growth via financial development in SSA, this paper determines the direct relationship between inflation and finance in the region. This is particularly important given that the region exhibits a low level of financial sector development as well as relatively high inflation compared to other regions (see Figure 1). Therefore, there is a need for tailor-made policy guidance. The paper considers a sample period associated with a host of external shocks, which include the 2008 global financial crisis, unstable exchange rates, and volatile commodity prices, which have influenced macroeconomic behaviour. The paper also goes an extra mile by considering a robust methodological approach at the frontier of the discipline: the dynamic panel threshold technique suggested by Seo et al. (2019) and Seo and Shin (2016), which considers the threshold to be endogenously determined along with other possible endogenous explanatory variables. In our model specification, inflation is endogenously determined in the model specification with other macroeconomic variables, justifying the need to use this approach in order to ensure the most effective results.

The remainder of this paper begins with some stylised facts on the inflation and financial development in Section 2, followed by a literature review in section 3. Section 4 provides the methodology and data. Sections 5 and 6 present the results and study conclusion, respectively.

2. INFLATION AND FINANCIAL DEVELOPMENT: STYLISED FACTS

Figure 1 shows regional average levels of inflation and financial development over the period 2000–2016. The SSA exhibits the least-developed financial sector (with the exception of low-income countries) as reflected by the credit-to-the-private-sector ratio, deposit money bank assets as a ratio of gross domestic product (GDP), and liquidity liability as a ratio of GDP. This implies a fragile financial system in SSA, which is worrying, as it can lead to destructive economic consequences for the progress of the region. SSA also had the highest level of inflation, measured by annual growth in the consumer price index (CPI). This implies high economic risk, which scares away potential investment and deters progress of the existing economic players in the region. As such, the region risks stagnant or declining economic growth in relation to other regions with better macroeconomic environments.

Figure 1: Average inflation and financial development by region, 2000–2016

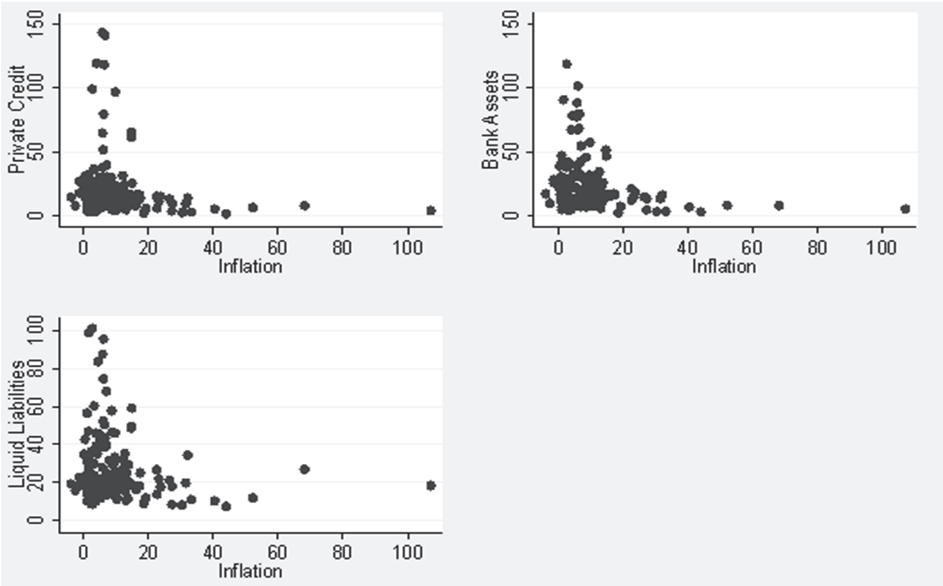


Source: Authors' computation based on World Development Indicators and Financial Development and Structure Dataset (2019)

Figure 2 shows scatter plots for inflation and financial development in SSA, averaged over the period 1982–2016. There is evidence of a non-linear relationship between the variables. This supports the idea that the relationship between inflation and financial development is complex and cannot be effectively analysed using general linear regressions. A positive correlation can be seen in lower inflation levels, while a negative relationship is also clear at higher inflation levels. This is a sure reflection of the existence of a non-linear relationship and hence the need to keep inflation levels below a certain level to ensure stability and development in the financial sector.

The individual countries with the lowest annual inflation (measured by consumer price index percentage) averaged over the period 2007–2016 are Zimbabwe (1.10%), Senegal (1.76%), Niger (1.87%), Burkina Faso (1.98%), Cabo Verde (2.13%), Côte d'Ivoire (2.17%), and Mali (2.24%). The slump in commodity prices, mainly for oil and food, has been the main indication of the subdued inflation in African countries, including Zimbabwe.

Figure 2: Financial development and inflation scatter plot for SSA, 1982–2016.



Source: Authors' own computations based on dataset from World Development Indicators and Financial Development and Structure (2019

Countries with relatively high annual inflation over the same period (2007–2016) are South Sudan (66.60%), Sudan (21.20%), Ethiopia (16.85%), Malawi (15.60%), Guinea (13.59%), Angola (13.54%), and Ghana (13.48%). The soaring inflation level in South Sudan can be attributed to civil war since 2013, while in Sudan inflation is a product of the economic challenges resulting from losing three-quarters of its oil revenue to South Sudan since 2011. In Angola the inflation has been driven by the government's effort to reduce the parallel and official exchange rates through devaluation of the kwanza, and a slump in commodity prices. Food inflation, in the light of maize shortages and the need for food aid, was the main driver of inflation in Malawi. Other major economies in the region over the same period (2007–2016) reported an average inflation rate of 10.65% for Nigeria, 8.91% for Tanzania, 9.81% for Kenya, 6.28% for Namibia, and 6.13% for South Africa.

In the same region over the same period, countries with highly developed financial sectors, measured by the private credit ratio provided by deposit money banks in the region, include Liberia (369.20%), Mauritius (89.33%), South Africa (68.50%), Cabo Verde (57.36%), Namibia (47.29%), Sao Tome and Principe (28.99%), Senegal (1.76%), Botswana (27.21%), and Kenya (27.19%). The countries with the least developed financial sectors over the same period are South Sudan (1.23%), the Democratic Republic of the Congo (4.12%), Chad (4.66%), Sierra Leone (5.34%), Guinea (5.50%), Equatorial Guinea (7.69%), the Republic of the Congo (7.86%), and Guinea-Bissau (7.92%). In the Republic of the Congo the depressed financial infrastructure which resulted in limited borrowing options can be attributed to depressed financial development. Guinea-Bissau is also associated with political instability, which led to the suspension of donor flows in 2016, while the weak financial development in South Sudan can be attributed to civil war since 2013, which jeopardised the major economic activities.

3. EMPIRICAL LITERATURE REVIEW

The literature on the relationship between inflation and financial development can be categorised into non-linear and linear studies. In the non-linear studies various thresholds have been established. Boyd et al. (2001), on a sample of between 65 and 97 countries (depending on data availability), establish a threshold of 15% by implying that economies with higher than the threshold inflation experience a discrete drop in financial sector development. Khan et al. (2006), on the other hand, arrive at an inflation threshold of about 3% to 6% for 168 industrialised and developing economies. They show that in a regime with inflation below the threshold the impact of inflation on financial markets conditions is insignificant and/or slightly positive, as it varies with the indicators used. A study by Tinoco-Zermeno et al. (2018) on 84 industrialised and less-industrialised countries finds a negative and non-linear impact of inflation on financial variables. The findings are particularly significant in the full sample and in developing economies, while they are not significant in developed economies.

In a study on Ghana, Abbey (2012) finds evidence for the existence of an inflation threshold ranging between 11% and 16%, below which the negative impact of inflation on financial development is statistically significant. Conclusions above the threshold are uncertain, as most coefficients become insignificant. However, Naceur and Ghazouani (2005) find no evidence of a threshold relationship in 11

Middle Eastern and North African countries (MENA), but observe a linear negative relationship between inflation and financial development.

Among studies that pay attention solely to the linear relationship between inflation and financial development, Al-Nasser and Jackson (2012) observe a negative connection for 15 Latin American countries using the fixed generalised least squares model. Bittencourt (2011) also finds a strong negative relationship between inflation and financial development for 10 Brazilian provinces.

In a study of 87 developed and less-developed countries and using the pooled mean group estimator method, Kim and Lin (2010) find contrasting evidence on the inflation and financial development nexus: it is negative in the long run and positive in the short run. Surprisingly, Korkmaz (2015) finds no connection between financial development and inflation for 10 selected European countries using the fixed effects method. Inflation, however, is found to affect economic growth. Focusing on only time series studies using ARDL bounds testing and the error correction method, Wahid et al. (2011) find that high inflation trends inhibit financial sector performance in Bangladesh in both the short and long run. Almalki and Batayneh (2015) find the same result in Saudi Arabia using the same method. Tinoco-Zermeno et al. (2014) also use the ARDL bounds test to establish the long-run influence of inflation on private credit provided by banks and economic growth in Mexico. In Iran, again using ARDL, Aboutorabi (2012) finds evidence of the deterrent effect of a high inflation rate on financial development.

This study seeks to fill a gap in the literature by investigating the direct relationship between inflation and financial development in SSA, in order to provide guidance for policy in the region. The study also reflects on the dynamic panel threshold technique as suggested by Seo et al. (2019) and Seo and Shin (2016), which considers the threshold variable to be endogenously determined with other possible endogenous explanatory variables. In a model specification with macroeconomic variables, inflation is, by and large, endogenously determined, justifying the need to utilise Seo and Shin's (2016) approach, which takes this factor into consideration to ensure the most effective results. To ensure the robustness of the findings, the study also presents results from Hansen's (1999) non-dynamic threshold approach, which does not allow the inclusion of a lagged dependent variable while ignoring the endogeneity of the threshold variable.

4. DATA AND METHODOLOGY

4.1 Data and variable description

The study considers a sample of SSA countries over the period 1982–2016 with 5-year-averaged data. The 5-year non-overlapping data periods for 1982–2016 are 1982–1986, 1987–1991, 1992–1996, 1997–2001, 2002–2006, 2007–2011, and 2012–2016. The non-overlapping averaged data was utilised so as to observe the long-run dynamics, which are not prone to temporary jumps, as is the case with annual data. The choice of 5-year averaged data does not negatively influence the effectiveness of the main results but instead provides a long-run interpretation that is not affected by short-term fluctuations in business cycles. For example, around the year 2008 there was a drastic change in the general flow of annual data, creating outliers in observations and distorting the long-run interpretation of the findings. Averaged data avoids a biased long-run data interpretation following the global financial crisis, which intensified in 2008, and proves to be superior to pure annual data.

Due to the need for strongly balanced data for the execution of the threshold approaches, only 23 countries are considered: Burundi, Burkina Faso, Botswana, Cameroon, Central African Republic, Chad, Republic of the Congo, Côte d'Ivoire, Gabon, Gambia, Ghana, Kenya, Madagascar, Malawi, Mauritius, Niger, Nigeria, Rwanda, Senegal, Seychelles, South Africa, Sudan, and Togo.

The financial development indicators used follow Bandura (2022), Kim and Lin (2010), Boyd et al. (2001), and Beck et al. (2000), and are private sector credit provided by deposit money banks as a ratio of GDP ('private credit'), liquid liabilities as a ratio of GDP ('liquidity liabilities'), and deposit money bank assets as a ratio of GDP ('bank assets'). They were all sourced from financial development and structure databases. The measure of credit distribution is the ratio of private credit by deposit money banks to GDP. The other financial development indicators used are the liquidity liabilities ratio and deposit money bank assets ratio, which represent the size of the financial sector. The liquidity liabilities of the financial sector as a ratio of GDP consist of the currency, demand,

and interest-bearing liabilities of the financial intermediaries (banks and non-banks) relative to the economy. The deposit money bank assets ratio is the total assets of money deposited in commercial and other deposit-taking banks as a ratio of GDP. Inflation is measured by the consumer price index (CPI). Annual growth is also a control variable used in the study and is expected to have an inverse relationship with financial development. The indicator is also utilised by Tinoco-Zermeno et al. (2014) and Zaman et al. (2010). Inflation rates and all the control variables were obtained from world development indicators.

Concerning the control variables, the actual value of secondary enrolment (% gross) is expected to positively contribute to financial development. The variable has been popularised in the literature by studies including Al-Nasser and Jackson (2012) and Naceur and Ghazouani (2005). Initial income as derived from GDP per capita (constant 2010 USD) is included in the study following Ductor and Grechyna (2015), Kim and Lin (2010), and Boyd et al. (2001). It is expected to have a positive impact on financial development. The actual value of real GDP per capita is also used, given that the panel threshold method only allows time-varying variables (Ehigiamusoe, et al., 2018; Almalki et al., 2015; Wahid et al., 2011; Khan et al., 2006; Naceur & Ghazouani, 2005). Trade as a ratio of GDP is used to account for external shocks. This variable is expected to have either a positive or negative relationship with financial development (Kim et al., 2012; Kim & Lin, 2010; Rajan & Zingales, 2003). Government consumption as a ratio of GDP is expected to be positively related to financial development. This variable is also used by Bittencourt (2011), Kim and Lin (2010) and Boyd et al. (2001). Gross fixed capital formation as a ratio of GDP is also expected to have a positive relationship with financial development. The variable is also utilised by Raheem and Oyinlola (2015) and Tinoco-Zermeno et al. (2014).

4.2 Econometric technique

The study utilises the pure cross-section method as baseline regressions for countries in SSA. The pure cross-section model, following Boyd et al. (2001), takes the form:

$$FD_i = \varphi + aInf_i + bInfd_i + cInfd_i * Inf_i + dw_i + e_i \tag{1}$$

where FD represents the financial development indicator, Inf is the inflation rate, and $Infd$ is the inflation dummy with 1 for inflation above the threshold and 0 if below the threshold. If inflation is greater than the threshold then the coefficient of inflation is $a+c$ and when it is below it is a; if inflation is greater than the threshold the intercept is $\varphi + c$, or φ if less than the threshold. w represents the vector of control variables, e is the error term, and i represents country.

The study uses the static and dynamic panel approaches as developed by Hansen (1999) and Seo and Shin (2016) to exploit the time series dimension of the data and dispense with the econometric problems associated with cross-section regressions, such as heterogeneity of groups. Hansen (1999) first developed a non-dynamic panel with individual specific effects, using least squares estimation of the threshold regression slopes using fixed effects transformation. To ensure the construction of a confidence interval and to test the hypotheses, the technique established a non-standard asymptotic theory of inference. The method also considers regression models that require explanatory variables to be endogenous and an exogenous threshold variable. The two-stage least squares estimator of the threshold parameter and a generalised method of moments estimator of the slope parameters was developed. As explained in Bandura (2022), the structural equation as proposed by Hansen (1999) takes the following form:

$$y_{it} = u_i + \beta_1^{'} x_{it} I(q_{it} \leq \gamma) + \beta_2^{'} x_{it} I(q_{it} > \gamma) + \varepsilon_{it} \tag{2}$$

Where $I(.)$ is the indicator function, an intuitive way of writing (2) is

$$y_{it} = \begin{cases} u_i + \beta_1^{'} x_{it} + \varepsilon_{it,}\ q_{it} \leq \gamma \\ u_i + \beta_2^{'} x_{it} + \varepsilon_{it}, q_{it} > \gamma \end{cases}$$

Another compact representation of *2* is to set

$$x_{it}(y) = \begin{pmatrix} x_{it} I(q_{it} \leq \gamma) \\ x_{it} I(q_{it} > \gamma) \end{pmatrix}$$

and $\beta = (\beta_1^{'}\ \beta_2^{'})'$ so that 2 equals;

$$y_{it} = u_i + \beta' x_{it}(\gamma) + \varepsilon_{it} \tag{3}$$

where y_{it} is the dependent variable, which in this study is financial development, and q_{it} is the time-varying threshold variable, which is inflation. β_1 and β_2 are the regression slope coefficients and are associated with regimes 1 and 2, respectively; γ is the threshold parameter; $I(.)$ is the indicator function; x_{it} represents the time-varying regressors; ε_{it} is the error term with mean zero and finite variance. One traditional method to eliminate the individual effect u_i is to remove individual-specific means. While straightforward in linear models, the non-linear specification (2) calls for a more special treatment. Note that taking averages of (2) over the time index *t* produces

$$\overline{y}_{it} = u_i + \beta_1' \overline{x}_{it}(\gamma) + \overline{\varepsilon}_{it} \tag{4}$$

Taking the difference between (3) and (4) yields

$$y_{it}^* = u_i + \beta_1' x_{it}^*(\gamma) + \varepsilon_{it}^*$$

where $y_{it}^* = y_{it} - y_{it}^*$

$$x_{it}^*(\gamma) = x_{it}(\gamma) - \overline{x}_{it}(\gamma)$$

$$\varepsilon_{it}^* = \varepsilon_{it} - \overline{\varepsilon}_{it}$$

Under the null hypothesis of no threshold, the model is

$$y_{it} = u_i + \beta_1' x_{it} + \varepsilon_{it}$$

After fixed-effects transformation

$$y_{it}^* = u_i + \beta_1' x_{it}^* + \varepsilon_{it}^*$$

The regression parameter β_1 is estimated by OLS, yielding $\tilde{\beta}_1$, residuals $\tilde{\varepsilon}_{it}^*$, and sum of squared errors $\tilde{\varepsilon}_{it}^{*'} \tilde{\varepsilon}_{it}^*$.

This study also takes advantage of the most recent user-friendly Stata command by Seo et al. (2019) for the panel threshold, which allows for the endogenous threshold variable and other advanced components. Seo and Shin (2016) advance the work of Hansen (1999), where endogenous variables for both the threshold variable and regressors are allowed on non-linear asymmetric dynamics and cross-sectional heterogeneity. They came up with first-differenced two-step least squares and first-differenced GMM methods, depending on whether or not the threshold variable is strictly exogenous. The first-differenced two-step least squares is ideal for a strictly exogenous threshold variable, while the first-differenced GMM is best with an endogenous threshold variable. As Seo and Shin (2016) provide the asymptotic distribution for both estimators, they also use a bootstrap-based test to check if there is any threshold effect, and an exogeneity test of the threshold variable. The approach takes the following form:

$$y_{it} = \left(1, x_{it}^{'}\right)\phi_1 I\left(q_{it} \le \gamma\right) + \left(1, x_{it}^{'}\right)\phi_2 I\left(q_{it} > \gamma\right) + \varepsilon_{it}, i = 1, \ldots, n; t = 1, \ldots, T, \quad (5)$$

where y_{it} is a scalar stochastic variable of interest, x_{it} is the $k_1 \times 1$ vector of time-varying regressors (may include lag of y_{it}), $I(.)$ is an indicator function, and q_{it} is the transition variables. The threshold parameter is represented by γ and the slope parameters associated with different regimes are ϕ_1 and ϕ_2. The regression error ε_{it} consists of the error components:

$$\varepsilon_{it} = \alpha_i + v_{it}$$

where α_i is an unobserved individual fixed effect and v_{it} is a zero mean idiosyncratic random disturbance. In particular, v_{it} is assumed to be a martingale difference sequence for the expositional simplicity,

$$E\left(v_{it} | \mathcal{F}_{t-1}\right) = 0$$

Where $\mathcal{F}_t$ is a natural filtration at time *t*, it should be noted that there is no assumption that x_{it} or q_{it} are to be measurable with respect to $\mathcal{F}_{t-1}$, thus allowing endogeneity in both the regressors, x_{it}, and the threshold variable, q_{it}.

5. EMPIRICAL RESULTS AND DISCUSSION

Table 1 shows that the averages reported for financial development range between 16.82% and 26.38% of GDP. These are much lower than found by Law et al. (2018) as the enabling percentage level of financial development for economic growth. Given the depressed level of financial development in the region, the result reflects the need to analyse the factors that determine development in the financial sector, as it is a prerequisite for the success of any economy. Inflation, on the other hand, shows a mean rate of 10%, which is higher than the acceptable level of inflation to facilitate development in the region. Averaged over the period 1982–2016, the low levels of financial development and the background of high inflation rates might signal a close negative relationship between the variables. As such, this study is key as it tries to identify the connection between the critical economic variables.

Table 1: Summary statistics

	Unit of measurement	Obs	Mean	Std, dev	Min	Max
Private Credit	Private credit by deposit money banks and other financial institutions (% of GDP)	161	19.25	23.32	1.23	143.37
Private credit banks	Private credit by deposit money banks (% of GDP)	161	16.82	15.34	1.23	99.17
Liquid liabilities	Liquid liabilities (% of GDP)	161	26.38	17.39	7.36	100.98
Bank assets	Deposit money bank assets (% of GDP)	161	22.63	19.79	2.17	118.68
Inflation	Percentage	161	9.72	12.54	-3.90	107.12
Trade	Total exports and imports (% of GDP)	161	67.18	33.26	13.80	206.00
GDP per capita	Constant 2010 US$	161	2264.02	2971.88	224.05	12825.14
Initial income	GDP per capita (constant 2010 US$)	161	1945.79	2690.20	294.91	12172.55
Sec Enrolment	School enrolment, secondary (% gross)	140	36.05	25.64	3.44	113.87
GFCF	Gross fixed capital formation (% of GDP)	161	20.02	8.59	3.70	64.40
GOV	General government final consumption expenditure (% of GDP)	161	14.67	6.39	1.39	40.90

Note: The table is based on 5-year averaged data from 1982–2016 with the exception of initial income, which is invariant (averaged between 1982 and 1984) for 23 SSA countries.

Source: Authors' computation

Table 2: Correlation matrix

	Private credit	Private credit banks	Liquid liabilities	Bank assets	Inflation	Trade	GDP per capita	Initial income	Sec Enrolment	GFCF	GOV
Private Credit	1.000	0.9203	0.5872	0.8138	-0.1475	0.1420	0.4433	0.3601	0.5695	0.0267	0.1932
Private credit banks		1.000	0.7397	0.8989	-0.1956	0.2569	0.4492	0.2903	0.5632	0.0657	0.1698
Liquid liabilities			1.000	0.9126	-0.2222	0.5663	0.5781	0.2849	0.6860	0.2331	0.3908
Bank assets				1.000	-0.2207	0.4156	0.6042	0.3792	0.7042	0.1637	0.3560
Inflation					1.000	-0.1225	-0.1683	-0.1843	-0.0416	-0.0739	-0.1655
Trade						1.000	0.5895	0.3909	0.4914	0.4495	0.3897
GDP per capita							1.000	0.8888	0.7300	0.4006	0.4543
Initial income								1.000	0.5463	0.3284	0.3321
Sec Enrolment									1.000	0.3337	0.3984
GFCF										1.000	0.0872
GOV											1.000

On the other hand, the much-hypothesised negative correlation between inflation and all the financial development indicators is also supported in Table 2. Liquidity liabilities are the most affected, followed by bank assets. Inflation also shows a negative relationship with the rest of the control variables used in the study. These control variables reflect the economies' macroeconomic stability. The control variables used are trade openness, real GDP per capita, initial income, secondary school enrolment, gross fixed capital formation, and government consumption. As such, it can be concluded that inflation is a threat to economic development. Table 3 shows the pure cross-section results.

The pure cross-section results in Table 3 show evidence of both a linear and a non-linear relationship between inflation and private credit. The increase in the inflation threshold (7% to 8%) indicates an increasing negative impact of inflation on private credit, which concurs with observations by Boyd et al. (2001). A significant negative impact of inflation on private credit is observed for an inflation level above the 7% and 8% thresholds. The negative coefficients are –0.4205 and –0.4819 respectively, obtained by summing *a* and *c* from the cross-section equation. There is, however, evidence of a positive impact of inflation on private credit at an inflation level below the threshold by interpreting only the coefficient of inflation (*a*) from the pure cross-section equation. Besides, the positive impact of inflation at 7% is 3.2492, which is bigger than 1.8103, obtained at the 8% inflation threshold and hence evidence of a stronger positive impact at lower inflation levels than at the higher inflation levels. Even though the same results can be concluded for liquid liabilities and bank assets, the majority of the coefficients are statistically insignificant.

Table 3: Pure cross-section regressions

Threshold regressions											
Variable	Inflation	InfD	InfD*inf	INV-INF	GOV	Trade	Initial income	Sec enrolment	Constant	R^2	Obs
(V) Threshold regression with 6% inflation threshold (robust)											
Private credit	2.1231 (1.536)	13.2791** (5.892)	−2.8686* (1.528)		−0.0330 (0.291)	−0.1590* (0.087)	0.0010 (0.001)	0.5675*** (0.156)	4.2404 (4.908)	0.37	140
Liquid liabilities	−0.1041 (0.679)	3.8920 (2.980)	−0.4585 (0.670)		0.2053 (0.251)	0.1602*** (0.044)	−0.0016*** (0.000)	0.4406*** (0.071)	1.4740 (4.111)	0.59	140
Bank assets	0.7245 (0.895)	5.4215 (3.954)	−1.3336 (0.882)		0.1910 (0.282)	0.0361 (0.054)	−0.0005 (0.001)	0.5315*** (0.103)	−0.0610 (4.663)	0.52	140
(W) Threshold regression with 7% inflation threshold (robust)											
Private credit	3.2492** (1.373)	8.2173 (5.247)	−3.6697*** (1.365)		0.0687 (0.288)	−0.1593* (0.084)	0.0012 (0.001)	0.5249*** (0.145)	2.0803 (5.045)	0.43	140
Liquid liabilities	0.6385 (0.553)	1.9482 (3.047)	−1.0564* (0.554)		0.2431 (0.246)	0.1589*** (0.045)	−0.0015*** (0.000)	0.4240*** (0.070)	0.3037 (4.055)	0.61	140
Bank assets	1.4538* (0.774)	1.9965 (3.857)	−1.8468 (0.770)		0.2566 (0.272)	0.0357 (0.053)	−0.0003 (0.001)	0.5040*** (0.099)	−1.4023 (4.601)	0.57	140
(X) Threshold regression with 8% inflation threshold (robust)											
Private credit	1.8103* (1.016)	6.4721 (4.971)	−2.2922** (1.018)		−0.0403 (0.290)	−0.1529* (0.086)	0.0009 (0.001)	0.5652*** (0.155)	5.2741 (4.485)	0.41	140
Liquid liabilities	0.3743 (0.483)	0.6037 (2.956)	−0.7618 (0.491)		0.2152 (0.248)	0.1610*** (0.045)	−0.0016*** (0.000)	0.4327*** (0.069)	0.9994 (4.050)	0.61	140
Bank assets	0.6421 (0.651)	0.6208 (3.760)	−1.0539 (0.655)		0.1925 (0.280)	0.0396 (0.054)	−0.0005 (0.000)	0.5272*** (0.100)	0.4462 (4.639)	0.55	140
(Y) Linear regression (robust)											
Private credit	−0.4199*** (0.149)				−0.1026 (0.295)	−0.1509* (0.090)	0.0008 (0.001)	0.6037*** (0.164)	11.703*** (3.661)	0.37	140
Liquid liabilities	−0.4364*** (0.090)				0.1917 (0.249)	0.1613*** (0.045)	−0.0017*** (0.000)	0.4475*** (0.071)	3.0366 (3.765)	0.61	140
Bank assets	−0.4891*** (0.107)				0.1597 (0.281)	0.0399 (0.054)	−0.0006 (0.001)	0.5479*** (0.102)	3.2508 (4.139)	0.54	140

Threshold regressions											
Variable	Inflation	InfD	InfD*inf	INV-INF	GOV	Trade	Initial income	Sec enrolment	Constant	R^2	Obs
(Z) Regression with inflation replaced by INV–INF (robust)											
Private credit				0.2483 (1.150)	−0.0369 (0.289)	−0.1447 (0.089)	0.0010 (0.001)	0.5843*** (0.161)	6.8925* (3.790)	0.36	140
Liquid liabilities				0.5110 (0.809)	0.2572 (0.251)	0.1671*** (0.046)	−0.0014*** (0.000)	0.4278*** (0.071)	−1.9946 (3.863)	0.57	140
Bank assets				0.2007 (0.607)	0.2372 (0.276)	0.0474 (0.055)	−0.0003 (0.001)	0.5251*** (0.101)	−2.3415 (4.228)	0.51	140

Note: robust standard errors in parentheses. *** indicates significance at 1%, ** significance at 5%, and * significance at 10%.

As shown in section Y of the results in Table 3, there is also strong evidence of a negative linear relationship between inflation and all the financial development indicators used in the study. However, section Z shows no evidence of a significant impact of inverse inflation (INV-INF) on any measure of financial development. The findings from this non-linear transformation are contrary to the expected non-linear nature of the relationship that is demonstrated by the threshold regressions. However, the coefficients are positive, supporting the theory underpinning the inverse relationship of inflation and financial development.

The control variables in Table 3 show mixed evidence. There is a positive and significant impact of human capital development (secondary school enrolment) on financial development, as expected from the literature. Government spending has an insignificant effect on financial development, as also found by Boyd et al. (2001). Initial income is largely insignificant, with a selective negative significant impact on financial development. The negative impact on financial development can be attributed to the convergence theory, which shows that countries with higher initial growth tend to have better financial systems, which in turn lead to a decreased rate of financial sector development as compared to countries with lower initial income. Trade openness is also found to have a negative impact on financial development, in line with findings by Kim et al. (2012) and Rajan and Zingales (2003).

Following Hansen (1999), Table 4 shows results for the threshold effect in non-dynamic panels for the inflation and financial development relationship. Generally, all the model specifications support the existence of a single threshold. The threshold estimator shows a level of inflation ranging between 6.39% and 6.52%, beyond which there would be a significant shift in the inflation and financial development relationship. There is evidence of a significant positive impact of inflation on financial development at low inflation rates (below the threshold), while the relationship turns negative (though not significant) at high inflation rates (above the threshold). The results endorse the idea that low inflation levels support positive development in the financial sector through the Philip's curve, while there is a high chance of endangering the same sector at inflation levels above the threshold by the way of credit market friction.

The estimated threshold of between 3% and 6% for global industrialised and developing economies, utilising the conditional least squares method, is slightly above that obtained by Khan et al. (2006). The difference can be related to the general observation that the inflation threshold is higher in less-industrialised economies such as the majority of countries in SSA, which have both higher productivity capacity and higher demand-pull inflation than developed economies. Generally, the positive impact of inflation below the observed threshold on financial development supports the existence of the short-run Philips curve. This implies that productivity is associated with a slight rise in prices that largely emanates from the demand side (demand-pull inflation). This follows an increased demand for both real and financial products. On the other hand, the inflation threshold obtained in this study is way lower than the range observed by Boyd et al. (2001), who provide evidence of a 15% threshold in developed and developing countries. This can be attributed to the different methodological approach used by Boyd et al. (2001), who did not use a specific panel threshold method.

The control variables in Table 4 show strong evidence of a positive and significant impact of gross fixed capital formation, GDP per capita, and government spending on financial development when inflation is below the threshold, and a negative relationship when inflation levels are above the threshold. These findings are in line with the expected results. However, there is a negative and significant impact on the relationship between trade and financial development in the region, which is not common, as the world (the African continent included) is pushing towards regional integration to ensure sustainable development. Kim et al. (2012) also observes these results in poorer countries and concludes that developing economies are not able to fully engage in gainful trading with more technologically advanced economies. It may also be explained by the need to ensure simultaneous opening of trade and capital flows so that the financial sector benefits from globalisation, as suggested by Rajan and Zingales (2003).

Table 4: Static panel threshold analysis

Financial development indicator	Private credit	Bank assets	Liquid liabilities
Threshold estimator $\widehat{\gamma}$	6.52%**	6.52%*	6.39%
95% confidence interval	[6.3941–6.5451]	[6.1325–6.5451]	[6.3890–6.5157]
Impact of inflation			
$\hat{\beta}_1$	1.5637*** (0.388)	1.0934*** (0.397)	0.7326* (0.400)
$\hat{\beta}_2$	–0.0569 (0.070)	–0.0914 (0.071)	–0.0282 (0.070)
Impact of covariates			
GOV	0.8982*** (0.217)	0.9141*** (0.222))	0.2533 (0.218)
GFCF	0.0380 (0.110)	0.1980* (0.112)	0.2430** (0.111)
Trade	–0.0865** (0.041)	–0.1275*** (0.042)	–0.0111 (0.0411)
GDP per capita	0.0073*** (0.001)	0.0079*** (0.001)	0.0051 (0.001)
Constant	–7.2467 (4.505)	–0.0914 (4.617)	6.1596 (4.541)
Observations	161	161	161
Countries	23	23	23

Note: Standard errors in parentheses. *** indicates significance at 1%, ** significance at 5%, and * significance at 10%. The results are based on the Hansen (1999) non-dynamic threshold approach.

Table 5 shows results for the threshold effect in dynamic panels for the inflation and financial development relationship, following Seo et al. (2019) and Seo and Shin (2016). The dynamic threshold approach allows the inclusion of the lagged dependent variable, which reduces the chances of model under-specifications, unlike the static method. The regressions are carried out restricting inflation as an endogenously determined variable. The results from the dynamic approach establish a lower threshold than the non-dynamic approach. The dynamic approach also estimates below- and above-threshold coefficients for all the variables in the model, including the control variables. Evidence of a significant threshold for all model specifications is reflected by rejecting the null hypothesis of linearity by bootstrap p-value for linearity test. The threshold estimator ranges between a 4.85% and 5.45% level of inflation with a significant shift in the inflation and financial development relationship. The threshold in the dynamic approach is slightly lower than in the non-dynamic method, which can be attributed to the many differences in the methodological framework, such as the inclusion of a lagged dependent variable and the restriction of inflation as endogenously determined in the dynamic models. Regarding the control variables in Table 5, there is strong evidence of a positive impact of trade and government spending on financial development when inflation is below the threshold and a negative relationship when inflation levels are above the threshold. These findings are in line with the expected results.

Generally, the obtained direct inflation threshold for financial development in SSA from both the dynamic and non-dynamic panel threshold techniques, which range between 4.85% and 6.62%, is below the indirect inflation threshold of 31% for economic growth through financial development found by Bandura (2022). The study uses almost the same dataset on the region. The findings imply a lag before the inflation threshold for growth is established via financial development, compared to the lower inflation threshold realised for the direct inflation and financial development nexus. Raheem and Oyinlola (2015) find an indirect inflation threshold of 15% for Nigeria and Cote D'Ivoire for the finance–growth nexus, which is also higher than the obtained direct inflation threshold in the current study on SSA. Even though the indirect inflation threshold they obtain for Ghana of between 5% and 10% is slightly lower, it is still higher than the range of 4.85% to 6.62% for the direct inflation–finance nexus.

Table 5: Dynamic panel threshold analysis following Seo et al. (2019) and Seo and Shin (2016)

Financial development indicator	Private credit	Bank assets	Liquid liabilities
Threshold estimator $\widehat{\gamma}$	5.45%	4.85%***	4.85%*
Bootstrapped p-value for linearity test	0	0	0
Impact of inflation			
$\hat{\phi}_1$	2.4950*** (0.793)	3.4637*** (1.182)	3.0630*** (0.599)
$\hat{\phi}_2$	–2.6776*** (0.771)	–3.6872*** (1.188)	–3.0838*** (0.545)
Impact of covariates below threshold			
Lagged dependent variable	0.8364*** (0.104)	0.1188 (0.227)	0.8587*** (0.137)
GOV	–0.1802 (0.505)	2.4344** (0.952)	1.5650*** (0.450)
Trade	0.1395*** (0.052)	0.1470*** (0.031)	0.0352 (0.035)
Impact of covariates above threshold			
Lagged dependent variable	0.0859** (0.038)	0.8123*** (0.243)	–0.0556 (0.125)
GOV	0.9696*** (0.357)	–2.3663** (0.952)	–1.1082*** (0.304)
Trade	–0.3278*** (0.058)	–0.2640 (0.067)	–0.0703* (0.042)
Constant	2.4950*** (5.017)	36.1108*** (7.651)	23.7759*** (3.837)
T	7	7	7
N	23	23	23

Note: standard errors in parentheses. *** indicates significance at 1%, ** significance at 5%, and * significance at 10%.

6. CONCLUSION

This paper analyses a sample of 23 countries in SSA over the period 1982–2016 using 5-year averaged data to determine the inflation threshold for financial development in the region. It uses a pure cross-section method and a panel threshold approach and finds strong evidence of a negative impact of inflation on financial development, which increases with a rise in inflation. The inflation threshold ranges between 4.85% and 6.62%, below which inflation has a positive impact on financial development, supporting the existence of a short-run Philips curve. There is, however, a negative relationship beyond the threshold. It is also worth noting that the obtained direct inflation threshold for financial development in SSA is lower that the established indirect impact of inflation on economic growth in the same region of 31% (Bandura, 2022). This shows that inflation can directly hinder financial development (and hence growth) at lower rates than was previously found for the indirect impact of financial development on economic growth.

It is therefore recommended that to ensure financial sector development in the region the authorities should adopt appropriate macroeconomic policies to keep inflation levels well below the observed threshold. Ensuring inflation levels below the range of 4.85% to 6.62% would encourage sustainable growth through financial sector development. The adoption of an inflation-targeting monetary policy could help to keep inflation levels in check at all times. Besides ensuring productivity growth, such a monetary policy could also help to contain inflation by matching demand and supply in the region.

The paper has some limitations. The threshold technique used requires a strongly balanced dataset: most sub-Saharan African countries have missing or incomplete data for the variables considered, which led to selection of only 23 countries with the required full set of data. Future studies could experiment with annual datasets, or different sample periods to further examine the robustness of the findings.

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THE IMPACT OF CORPORATE GOVERNANCE CHARACTERISTICS ON COMPANIES' FINANCIAL PERFORMANCE: EVIDENCE FROM ROMANIA

ABSTRACT: *This paper assesses the possibility of a relationship between corporate governance mechanisms, as independent variables, and firm performance measures, as dependent variables. The data was taken from the annual reports of a sample of 66 companies listed on Bucharest Stock Exchange in Premium and Standard categories during the period 2016–2020. The SPSS statistical program was used to run the multivariate linear regression model on the selected sample. Additional variables were used to control for leverage and size. The results of the study are mixed. Board size, board gender, and board meetings have a positive impact on a firm's performance, measured by both return on assets (ROA) and return on equity (ROE). CEO duality has a positive and significant impact on a firm's performance measured by ROA, while a negative and insignificant correlation was founded for ROE. Board independence has a negative and insignificant association with both firm performance measures. The results obtained can help companies to manage their corporate governance.*

KEY WORDS: *corporate governance, firm's performance, Romanian listed companies*

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1. INTRODUCTION

The link between corporate governance and firm performance has been widely debated in recent years, especially after the 2008 crisis. Multiple studies have analysed the link between the different characteristics of both corporate governance and a firm's performance and various authors have studied this correlation using different corporate governance mechanisms and measurements.[1] Lungu et al. (2020) and Mititean and Constantinescu (2020) suggest that the most common variables used to measure corporate governance are governance index, board size, board gender, chief executive officer (CEO) duality[2], and board independence, while for return on assets (ROA) and return on equity (ROE), Tobin's Q (TQ) ratio and market value are used to measure firm performance.

Previous studies have produced mixed results regarding the link between corporate governance and firm performance. Khatib and Nour (2021), Choi et al. (2020), and Khan et al. (2019) find that board size has a positive impact on firm performance, while Cheng (2008) and Al-Matari et al. (2012) find a negative relationship. Board independence is negatively associated with firm performance according to Koji et al. (2020) and Terjesen (2015) and positively associated according to Duru et al. (2016), Uribe-Bohorquez et al. (2018), and Tleubayev et al. (2020).

The aim of this research is to investigate the possible associations between corporate governance mechanisms and firm performance for a sample of 66 companies listed on Bucharest Stock Exchange in the Premium and Standard categories. This paper uses the quantitative method used by most researchers on emerging European countries, as suggested by Mititean and Constantinescu (2020). Thus, the paper intends to evaluate the impact of board size, board independence, board gender diversity, board meetings, and CEO duality on corporate performance, measured by a company's ROA and ROE.

1 For example Wang et al., 2017; Borlea et al., 2017; Rashid 2018; Choi et al., 2020; Ciftci et al., 2019; Duppati et al., 2019; Wijethilake & Ekanayake, 2019; Papangkorn et al., 2019; Liu & Jiang, 2020; Hsu et al., 2019; Pintea et al., 2020; Idris & Ousama, 2021; and Khatib & Nour 2021.

2 CEO duality occurs when the CEO serves both as chief executive officer and chairs the company board.

The obtained results are mixed. For instance, board size, board gender, and board meetings have a positive impact on firm performance as measured by ROA and ROE, while CEO duality and board independence have significant positive and negative impacts respectively on a firm's performance, as measured by ROA. The analysis was extended to cover 2019 and 2020 in order to investigate the impact of the COVID-19 pandemic on firm performance. The results suggest that board gender has a positive impact on firm performance in uncertain times for ROA while CEO duality negatively affect this relationship.

The rest of this paper is organised as follows: the second section presents the most recent literature on the subject, while the third section presents the research design and method, variables, and data sample. The fourth section presents the results of the descriptive statistics, Pearson correlation and Spearman rank correlation, and linear regression. The fifth section concludes.

2. LITERATURE REVIEW AND HYPOTHESES DEVELOPMENT

The relationship between corporate governance and performance has been widely debated at both the national and international level (Mititean & Constantinescu, 2020). An interesting study conducted by Lungu et al. (2020) shows that in emerging countries the most common metrics used by researchers from emerging economies to measure corporate governance are the board's independence, the corporate governance index, and the board's size, while performance is measured by ROA, ROE, and Tobin's Q ratio.

2.1 Board size and financial performance

Many empirical investigations of the relationship between board size and business performance produce mixed results. Wang et al. (2017) study the influence of board size on company performance in the Taiwanese hotel industry. Using a panel regression model on a sample of 448 observations during 64 quarters (1998–2013) and using the total number of directors on the company board as the metric for corporate governance and sales growth rates, and ROA, ROE, and TQ as performance metrics, the results show that a smaller board has a positive impact on company performance, while a bigger board has a negative impact. Choi et al. (2020) find that board size has a significant impact on financial performance in alcohol industry companies in the United States during the period 2003–2017. Merendino and Melville (2019) study this relationship for Italian listed companies during 2013 and 2015 and find a positive association between board size and firm performance. Ciftci et al. (2019) find a positive effect of board size on firm performance.

Borlea et al. (2017) study the relationship between board characteristics and firm performance for a sample of 55 Romanian non-financial companies listed on the Bucharest Stock Exchange in 2012 and find no statistically significant association. Duppati et al. (2019) find a negative association between board size and company performance for the non-financial companies listed on the SGX Mainboard in Singapore and the National Stock Exchange (NSE) in India during the period 2005–2015. Board size is found to have a positive effect on company performance during uncertain times in a study conducted by Khatib and Nour (2021). Hermuningsih et al. (2020) find that board size significantly improves firm performance.

Based on the results of previous studies, we conclude that studies on developed countries show a positive association between board size and firm performance, while for studies conducted on developing countries the results are mixed. Thus, our first hypothesis is:

H1: Board size has a positive impact on financial performance in Romanian listed companies.

2.2 Board independence and financial performance

Independent directors must be qualified for certain activities established by the board of directors and are not majority shareholders. Shan (2019) studies the impact of board independence and managerial ownership on firm performance using 9,302 firm-year observations for Australian listed companies in the period 2005–2015. The results indicate a negative effect of board independence on firm performance and/or vice versa. Merendino and Melville (2019) find a positive effect of board independence on firm performance. Liu et al. (2015) study the association between board independence and firm performance for a sample of 2,057 firms listed on the Shanghai and Shenzhen Stock Exchanges during 1999–2012, collecting a total of 16.999 firm-year observations. Measuring firm performance using ROA, ROE, and TQ, the results shows that independent directors have a positive effect on firm performance. Li and Roberts (2018) conduct a study of New Zeeland for the period 2004–2016 analysing the association between board independence and firm performance and find that board independence does not improve firm performance. Thus, the results are mixed, which may be explained by the different industries and years of observation.

Some studies are concerned with the relationship between board independence and firm performance in developing countries. Rashid (2018) analyses the impact of board independence on company performance in 135 firms listed on the Dhaka Stock Exchange during 2006–2011. The results suggest that board independence does not have a positive impact on firm performance. Borlea et al. (2017) also find no significant relationship between these variables. Khatib and Nour (2021) find a negative effect in uncertain times. Based on these results we develop our second hypothesis:

H2: Board independence has a positive impact on financial performance in Romanian listed companies.

2.3 Board meetings and financial performance

Frequent board meetings help managers to understand the main problems arising in their firms (Hanh et al. 2018). However, Hanh et al. (2018) examine the effect of board meetings on company performance using a sample of 94 companies listed on the Ho Chi Minh Stock Exchange during the period 2013–2015 and find that board meetings negatively affect company performance. Khatib and Nour (2021) also suggest a significant negative influence on firm performance.

Eluyela et al. (2018) study the impact of board meeting frequency on company performance in 15 deposit money banks on the Nigerian Stock Exchange during the period 2006–2011. The results shows that board meetings have a positive impact on company performance. Idris and Ousama (2021) examine board meetings and firm performance for a sample of 42 companies listed on the Qatar Stock Exchange (QSE) for the year 2018. Using two regression models (ROA and ROE), the results reveal that board meetings have a positive impact on firm independence. Our third hypothesis is based on these studies:

H3: Board meetings have a positive impact on financial performance in Romanian listed companies.

2.4 Board gender diversity and financial performance

Studies that look at gender differences in business have shown that women lead differently from men: they tend to mitigate conflict by being more collaborative (Bart and McQueen 2013, Gipson et al. 2017, Kirsch 2018*).* Papangkorn et al. (2019) examine the impact of female directors on firm performance using a sample of 16,156 firm-year observations for the years 2008 and 2009. Their study shows that gender diversity on a board has a positive impact on firm performance.

Duppati et al. (2019) analyse the influence of gender diversity on company performance in non-financial companies listed on the SGX Mainboard in Singapore and the National Stock Exchange (NSE) in India during the period 2005–2015. Using multiple regression analysis on 8,833 firm-year observations, the results reveal that gender diversity has a positive impact on company performance. Idris and Ousama (2021) find a positive relationship between gender diversity and financial performance. In these studies, board gender diversity is measured as the ratio of the total number of female directors on boards to the total number of directors on boards. Based on these results, our fourth hypothesis is developed:

H4: Board gender diversity has a positive impact on financial performance in Romanian listed companies.

2.5 CEO duality and financial performance

The impact of CEO duality on firm performance has been widely debated, with mixed results. In an analysis of Romanian and Bulgarian banks during 2005–2015, Onofrei et al. (2018) find that if CEO duality exists it has a negative and statistically significant impact on bank performance, while the absence of CEO duality has a positive impact. Wijethilake and Ekanayake (2019) apply multiple regression on a sample of 212 companies listed on the Colombo Stock Exchange in Sri Lanka for the year 2009. The results shows that the CEO not being the head of the executive board has a negative effect on firm performance, while CEO duality has a positive effect on company performance.

On the other hand, Hsu et al. (2019) find that CEO duality has statistically significant negative impacts on company performance, while Chang et al. (2018) find that CEO duality has a positive effect on company performance. Uppal (2020) analyses this relationship for the auto industry in India during 2011–2016. The results show that CEO duality can significantly impact this relationship. In these studies, CEO duality is a dummy variable that equals 1 if CEO and chairman are not separate roles and 0 otherwise. Based on the previous results, our fifth hypothesis is developed:

H5: CEO duality has a positive impact on financial performance in Romanian listed companies.

3. RESEARCH DESIGN AND METHODOLOGY

3.1 Data sample

The aim of this study is to examine the impact of corporate governance mechanisms on firm performance in an emerging country. The data for our study was collected from the annual reports of the Premium and Standard companies listed on the Bucharest Stock Exchange during the period 2016–2020. Financial companies are excluded because they are highly leveraged and subject to different regulations. Also excluded are companies without data for the whole period. The 10 industries included in our sample are listed in Table 1.

Table 1: Sample and industries

Industry	**No of companies**	**% of sample**
Accommodation and Food Service Activities	4	6%
Construction	4	6%
Manufacturing	41	62%
Mining and Quarrying	4	6%
Transportation and Storage	4	6%
Wholesale and Retail Trade	3	5%
Other	6	9%
Final sample	**66**	**100%**

Source: Authors' calculation

3.2 Variables

Details of the variables used in this research are summarised in Table 2. To have a holistic approach to the dependent variable measuring company performance, two financial indicators were considered: Return on Assets and Return on Equity. Because the focus of this study is financial performance, we choose accounting-based measures, as many other authors use this metric (Bachmann et al., 2019; Ciftci et al., 2019; Wang et al. 2017; Detthamrong et al., 2017; Gaur et al., 2015; Koji et al., 2020; Kyere and Ausloos, 2020; Lam et al., 2013; Mishra et al., 2020; Din et al., 2021).

Five independent variables were taken into consideration: board size, board independence, board meeting frequency, gender diversity, and CEO duality. Two control variables were included: leverage, calculated as the ratio of total debt to total assets, and firm size, calculated as the natural logarithm of total assets.

Table 2: Variables used in the linear regression model

Variable	Proxy	Type	Description	Referenced studies/research
Board size	BZ	I	Total number of board members	Arora and Sharma (2016); Bachmann et al. (2019); Christensen et al. (2010); Ciftci et al. (2019); Detthamrong et al. (2017); Gaur et al. (2015); Hamutyinei et al. (2015); Kılıç and Kuzey (2016)
Board independence	BI	I	The number of independent directors on the board divided by the total number of board members	Arora and Sharma (2016); Bachmann et al. (2019); Christensen et al. (2010); Ciftci et al. (2019); Wang et al. (2017); Shuaib et al. (2021); Detthamrong et al. (2017); Hamutyinei et al. (2015); Kılıç and Kuzey (2016)
Board meetings	BM	I	The number of board meetings held every year	Christensen et al. (2010); Hamutyinei et al. (2015); Koji et al. (2020); Papangkorn et al. (2019) and Khatib and Nour (2021)

Variable	Proxy	Type	Description	Referenced studies/research
Board gender diversity	BG	I	The ratio of the total number of female directors on the board to the total number of board members	Ciftci et al. (2019); Wang et al. (2017); Detthamrong et al. (2017); Kılıç and Kuzey (2016); Papangkorn et al. (2019); Khatib and Nour (2021); Duppati et al. (2019); Li and Chen (2018); Marinova et al. (2016)
CEO Duality	CEO	I	Dummy variable: equals 1 when CEO doubles as board chair and 0 otherwise	Wijethilake and Ekanayake (2019); Arora and Sharma (2016); Christensen et al. (2010); Ciftci et al. (2019); Shuaib et al. (2021); Detthamrong et al. (2017); Gaur et al. (2015) and Kyere and Ausloos (2020)
Return on Assets	ROA	D	The ratio of earnings before interest and taxes to total assets	Bachmann et al. (2019); Ciftci et al. (2019); Wang et al. (2017); Detthamrong et al. (2017); Gaur et al. (2015); Koji et al. (2020); Kyere and Ausloos (2020); Lam et al. (2013); Mishra et al. (2020) and Din et al (2021)
Return on Equity	ROE	D	Profit after tax as percentage of total equity	Wang et al. (2017); Detthamrong et al. (2017); Lam et al. (2013); Din et al (2021) and Khatib and Nour (2021)
Leverage	LV	C	Ratio of total debt to total assets	Akbar et al. (2016); Arora and Sharma (2016); Ciftci et al. (2019); Shuaib et al. (2021); Detthamrong et al. (2017); Hamutyinei et al. (2015); Kılıç and Kuzey (2016) and Kyere and Ausloos (2020)
Firm size	FZ	C	Natural logarithm of total assets	Kyere and Ausloos (2020); Mishra et al. (2020); Din et al (2021); Khatib and Nour (2021); Duppati et al. (2019) and Li and Chen (2018

Notes: I – Independent variable**; D** – dependent variable; **c** – Control variable;

3.3 Research method

Linear regression analysis was used to determine the relationship between corporate governance mechanisms and firm performance. Regression analysis has been widely used by other researchers (Ciftci et al., 2019; Wang et al., 2017; Detthamrong et al., 2017; Kılıç and Kuzey, 2016; Papangkorn et al., 2019; Khatib and Nour, 2021; Duppati et al., 2019; Li and Chen 2018; Marinova et al., 2016). The SPSS statistical program was used to run the regression model on the selected sample. The regression model used to analyse the influence of corporate governance mechanisms on company performance is expressed as follows:

$$FP_{it} = \beta_1 BZ_{it} + \beta_2 BIND_{it} + \beta_3 BM_{it} + \beta_4 BG_{it} + \beta_5 CEO_{it} + \beta_6 LV_{it} + \beta_5 FZ_{it} + \varepsilon_{it}$$

where *FP* is firm performance, which subsequently takes the value of the return on assets (ROA) and return on equity (ROE), *BZ* is board size, *BIND* is board independence, *BM* is board meeting, *BG* is board gender diversity, *CEO* is CEO duality, *LV* is leverage, *FZ* is firm size, β_{1-5} are regression coefficients, and ε_{it} is the error term.

4. RESULTS

4.1 Descriptive statistics and correlation matrix

Table 3 presents a descriptive analysis of corporate governance characteristics, company performance, and other firm-level control variables for the Bucharest Stock Exchange-listed firms. The sampled firms show a mean ROA value of 0.03 and mean ROE value of 0.11. The minimum ROE is –16.09 and the maximum value is 14.08, which can be translated as a difference between companies.

The mean of board size is 4.54, and 46% of the boards are independent. Twenty-eight per cent of the CEOs are also board chairs, and the mean of board meetings is 15.05 per year. The gender diversity average of Bucharest Stock Exchange listed firms is 0.19, ranging between 0.00% and 100%. This low percentage of gender diversity on boards suggests that the companies listed on the BSE have little female representation.

Table 4 reports the Pearson (below the diagonal) and Spearman (above the diagonal) correlation matrix for all the variables. Board size and CEO duality are positively correlated with ROA at the level of 0.001 and 0.05 respectively, while board independence in negatively correlated with ROA at the 0.05 level.

Table 3: Descriptive Statistics of Variables

Variable	N	Min	Max	Mean	Std. Dev.	Variance	Skewness		Kurtosis	
							S	SE	S	SE
BZ	325	1.00	11.00	4.54	1.78	3.17	0.31	0.14	0.54	0.27
BI	225	0.00	1.00	0.46	0.30	0.09	0.49	0.16	−0.74	0.32
BG	312	0.00	1.00	0.19	0.22	0.05	1.16	0.14	1.42	0.28
BM	214	0.00	60.00	15.05	11.43	130.55	1.64	0.17	2.56	0.33
CEO	319	0.00	1.00	0.28	0.45	0.20	0.97	0.14	−1.06	0.27
ROA	325	−1.11	2.09	0.03	0.17	0.03	3.91	0.14	69.96	0.27
ROE	325	−16.09	14.08	0.11	1.32	1.75	−1.62	0.14	109.31	0.27
LV	325	−2.05	5.04	0.52	0.69	0.47	3.36	0.14	16.50	0.27
FZ	325	12.04	24.56	19.07	1.96	3.86	−0.15	0.14	1.87	0.27

Notes: S – Statistic**; SE** – Std. Error

Source: Authors' calculation

Table 4: Pearson/Spearman correlation matrix

Variable	BZ	BI	BG	BM	CEO	LV	FZ	ROA	ROE
BZ	1	−.266**	0.05	.293**	−0.014	−.255**	.496**	.350**	.115*
BI	−0.063	1	−0.064	.148*	−.183**	0.088	−0.034	−.139*	−0.106
BG	−.160*	−0.022	1	−0.034	0.022	−.160**	0.001	0.036	−.129*
BM	.259**	.213**	−0.094	1	−.194**	−0.017	.362**	0.104	−0.002
CEO	−0.062	−.186**	0.042	−.205**	1	0.063	−0.04	.141**	.144**
ROA	.231**	−.147*	0.074	0.125	.147*	−.223**	.197**	1	.485**
ROE	0.013	−0.069	0.117	−0.020	0.036	.304**	.098*	−0.061	1
LV	−.143*	−0.043	−0.081	−0.115	0.068	1	0.035	−.509**	.233**
FZ	.475**	0.083	−.212**	.428**	−.237**	0.000	1	0.053	−0.056
VIF	1.37	1.09	1.06	1.31	1.11	1.06	1.59	–	–
Tolerance	0.73	0.92	0.94	0.76	0.90	0.95	0.63	–	–

Notes: In the above table, Pearson (Spearman) correlations are presented below (above) the diagonal of the matrix. *Correlation is significant at the 0.05 level **Correlation is significant at the 0.01 level (1-tailed).

Source: Authors' calculation

No correlation was identified for the dependent variable ROE according to the Pearson correlation matrix. However, the Spearman correlation matrix positively correlates board size with ROA at the 0.001 level and with ROE at the 0.05 level. Board independence is negatively correlated with ROA, while board gender has the same correlation with ROE. CEO duality is positively correlated at the 0.001 level with both ROE and ROA.

We use the variance inflation factor (VIF) to check for potential multicollinearity issues. The results shows that the VIFs for the independent variables are below 10 (the range is between 1.06 and 1.59) and the tolerance range of between 0.63 and 0.95 is above 0.1, which means that there is no multicollinearity, in accordance with Shan (2015) and Wang et al. (2019).

4.2 Regression results and discussion of research hypotheses

Table 5 presents the relationship between corporate governance mechanisms, represented by board size (BZ), board independence (BI), board gender (BG), board meetings (BM), and CEO duality (CEO), and firm performance, represented by ROA and ROE. The coefficient of board size is positive for both ROA and ROE, but the results are insignificant (Sig. >0.05), thus supporting H1. This means that larger boards have a positive impact on firm performance. Our results are consistent with studies by Khatib and Nour (2021), Choi et al. (2020) and Khan et al. (2019), but contradict Cheng (2008) and Al-Matari et al. (2012).

The percentage of board independence is negative and significant for ROA (Sig. <0.05) and negative and insignificant for ROE; thus H2 is not supported. The negative association between board independence and firm performance is in accordance with prior studies by Koji et al. (2020) and Terjesen (2015) but contrary to the results of Duru et al. (2016), Uribe-Bohorquez et al. (2018), and Tleubayev et al. (2020). The results suggest that for BSE-listed firms the percentage of independent board members does not increase firm performance.

Table 5: The impact of corporate governance mechanisms on firm performance

Dependent	ROA		ROE	
Independent	**Coefficient**	**Sig.**	**Coefficient**	**Sig.**
(Constant)	0.053	0.424	0.123	0.633
BZ	0.008	0.058	0.019	0.258
BI	–0.057	0.029	–0.066	0.514
BG	0.029	0.379	0.229	0.076
BM	0.001	0.162	0.001	0.591
CEO	0.041	0.007	–0.002	0.974
LV	–0.167	0.000	0.293	0.001
FZ	0.000	0.905	–0.013	0.339
F statistic	12.391		2.174	
Durbin-Watson	2.131		1.813	
Adjusted R-square	0.349		0.046	
ANOVA Sig	<.001[b]		0.039[b]	
N	325			

Source: Authors' calculation

Regarding H3, gender diversity has a positive correlation with ROA and ROE but it is insignificant (Sig. >0.05), which means that the percentage of women on the board increases the level of firm performance but not significantly. Thus, H3 is supported, in accordance with prior studies by Li and Chen (2018), Duppati et al. (2019), Bin Khidmat et al. (2020), and Đặng et al. (2020).

Board meetings have a positive impact on ROA and ROE but it is insignificant, which suggests that H4 is supported. Our results are in accordance with Eluyela et al. (2018) but contrary to Hanh et al. (2018).

Finally, CEO duality has a positive impact on firm performance for ROA at the 0.05 significance level and an insignificant positive impact on ROE, which means that H5 is supported. Our result are contrary to Wijethilake and Ekanayake (2019), Hsu et al .(2019), and Tang (2017), which suggests that if the same person is both chair of the board and chief executive officer the effect on firm performance is negative.

4.3 Additional analysis

We conduct an additional test for the years 2019 and 2020 to see if the COVID-19 pandemic impacts the link between corporate governance mechanisms and firm performance. We split our sample into 2019 and 2020 subsamples for both ROA and ROE firm characteristics, as shown in Table 6.

The results suggest that in 2020, during times of uncertainty, board gender diversity had a positive impact on firm performance characterized by ROA, while in the previous year it had an inverse association that was negative and insignificant. CEO duality had a negative but insignificant impact during the crises but a positive impact on firm performance in 2019. Board size had a positive impact during uncertain times in 2020 but a negative impact in 2019, both insignificant, which means that larger boards helped to improve firm performance measured by ROE during the crisis period.

Table 6: The impact of corporate governance mechanisms on firm performance: Year subsample

	ROA				ROE			
	2020		2019		2020		2019	
	B	Sig.	B	Sig.	B	Sig.	B	Sig.
(Constant)	–0.032	0.785	0.211	0.183	–0.228	0.692	0.340	0.239
BZ	0.012	0.062	–0.004	0.628	0.021	0.494	–0.002	0.918
BI	–0.050	0.304	–0.080	0.177	–0.131	0.586	–0.080	0.454
BG	0.188	0.025	–0.022	0.742	0.576	0.153	0.115	0.349
BM	0.000	0.735	0.001	0.283	0.000	0.983	0.002	0.468
CEO	–0.008	0.804	0.075	0.021	0.206	0.177	0.019	0.739
LV	–0.069	0.148	–0.142	0.014	–0.395	0.096	0.292	0.007
FZ	0.001	0.896	–0.005	0.590	0.010	0.720	–0.018	0.252
F statistic	2.296		1.763		1.333		2.190	
Durbin-Watson	2.151		2.306		1.869		1.684	
Adjusted R-square	0.197		0.126		0.070		0.184	
ANOVA Sig	0.061[b]		0.132[b]		0.278[b]		0.064[b]	
N	32		38		32		38	

Source: Authors' calculation

Board meetings had a positive impact on ROA and ROE during both years. The percentage of women on boards and CEO duality had a positive impact on firm performance measured by ROE, while board independence had a negative impact for the BSE-listed firms for both periods, which means that if the boards had more independent directors, performance would not increase.

5. CONCLUSION

This study examines the impact of corporate governance mechanisms on firm performance in an Emerging European Country using descriptive statistics and regression analysis for a sample of 66 companies listed on the Bucharest Stock Exchange in Premium and Standard categories during the period 2016–2020. The corporate governance mechanisms used are board size, board independence, board meetings, board gender diversity, and CEO duality. Firm performance was measured by ROA and ROE, while firm size and leverage were control variables. The data sample was realised manually by taking the information from the annual reports of the listed companies.

The regression results suggest that board size, board gender, and board meetings have a positive impact on firm performance, measured by both ROA and ROE, supporting hypotheses H1, H3, and H4. These results are similar to the results of Khatib and Nour (2021), Choi et al. (2020), Khan et al. (2019), Li and Chen (2018), Duppati et al. (2019), Bin Khidmat et al. (2020), and Eluyela et al. (2018). H2, regarding to the positive impact of board independence on firm performance measured by ROE and ROA, is not supported; the results show a negative relationship between these characteristics. CEO duality is positively and significantly correlated with ROA and negatively correlated with ROE, while board independence is significant but negatively correlated with firm performance measured by ROA.

This study also investigates the impact of the COVID-19 pandemic on the relationship between corporate governance mechanisms and financial performance. After dividing our sample into the years 2020 and 2019 the regression results suggest that in uncertain times board gender has a positive and significant impact on firm performance measured by ROA, while a negative but insignificant relationship was identified for CEO duality in uncertain times. In addition, board size had a positive but insignificant impact on firm performance measured by ROE during uncertain times.

Like other studies conducted on emerging countries, our results were mixed. In our study on Romania, board size has a positive impact on both ROA and ROE, while Duppati et al. (2019) found a negative relationship for companies in Singapore. Our results partly agree with Borlea et al. (2017), who found board size to be insignificant. Board independence has a negative impact on ROA (significant) and ROE (insignificant), in accordance with the results of Rashid (2018), who studied this relationship for Bangladesh, and partly agreeing with Borlea et al. (2017), who found an insignificant impact on financial performance. Board meetings and board gender diversity have an insignificant positive effect on financial performance in our study, as Eluyela et al. (2018) found for board meetings in their study on Nigeria. This result partially agrees with the results of Duppati et al. (2019), who found a positive and significant impact. CEO duality has a positive impact on ROA (significant) and ROE (insignificant), unlike the results of Onofrei et al. (2018) who found a negative relationship for Bulgaria, and partially in accordance with Ekanayake (2019).

Our findings have important implications for companies, shareholders, regulators, and government because they suggest that companies and regulators should improve their reporting and establish new rules for corporate governance. This paper fulfils an identified need to study how corporate governance mechanisms can affect firm performance and contributes to the literature by offering new insights into the link between corporate governance and firm performance in an Emerging European Country.

Our study has some limitations. First, due to data limitations the number of governance variables was restricted. Second, some companies were not included in the study due to the unavailability of data for the chosen variables. Future research could be extended to more corporate governance mechanisms and more firm performance measurements. The sample could also be extended by taking one industry and collecting data for more Emerging European Countries in order to find possible new patterns.

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