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THE IMPACT OF US ELECTIONS ON THE DOLLAR'S EXCHANGE RATE

ABSTRACT: *This paper explores the effect of U.S. domestic politics on the behaviour of international currency markets. Specifically, for the first time in the literature, we gauge the impact of a divided government on the exchange rate volatility of five currencies: the Japanese yen, the Canadian dollar, the British pound, the Mexican peso, and the euro. At the same time, we control for the impact of political and macroeconomic factors. A GARCH methodology has been adopted for this objective, using weekly data from 2000 to 2021. The evidence suggests that the partisan and divided government variables significantly impact the conditional variance equation, whilst the observed reduced levels of exchange rate volatility during a Democrat presidency run counter to prior studies on partisanship. In addition, exchange rate volatility seems to increase one month before an election and during periods of divided government. Given the nascent evidence, we argue that U.S. politics are instrumental in affecting global financial markets.*

KEY WORDS: *International economics, exchange rates, political economy, GARCH modelling.*

JEL CLASSIFICATION: F31, G15, P16, C22.

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1. INTRODUCTION

Since the 1970s and the collapse of the Bretton Woods system, exchange rate volatility has been at the centre stage of academic research. Generally, most of the extant academic literature on the causes of exchange rate volatility focuses on financial and macroeconomic factors (see Brogaard et al., 2020; Vogiazas et al., 2019). However, Lobo and Tufte (1998), from a political economy perspective, were amongst the first to explore the channels through which politics can affect exchange rate volatility. The incorporation of electoral and partisan variables into the financial literature was also observed in studies by Alesina (1988), Alesina and Sachs (1988), and Alesina and Roubini (1992).

Ohmae (1995) argued that nations lose their ability to control and protect their exchange rates and therefore forfeit their roles as critical participants in the global economy. This narrative, which prevailed in the 1990s, reflects the economic and political landscape studied by most of the prior literature, such as Bachman (1992), Lobo and Tufte (1998), and Blomberg and Hess (1997).

Over the last decades, the global economic landscape has changed dramatically due to the growing complexity of international trade, multinationals' role, and the increased economic interdependence between countries. More recently, we have also observed the polarisation of U.S. politics, the rise of populism, and the growing resentment towards globalisation.

Intriguingly, the research in this area mainly focuses on electoral variables, but the results are conflicting or inconclusive. Therefore, we argue that partisan factors would be important given the ever-changing nature of partisan politics. The ideological composition of the Republican and Democrat parties in the U.S. is quite distinct from that of 30 years ago, as defined by the present societal and economic conditions. As most research on partisan effects took place more than 30 years ago, we feel that there is vital scope for a study looking at a more nuanced perspective of domestic politics and the role of international politics on exchange rate volatility. Given the unravelling importance of globalisation, such an approach will provide further insight into the fundamental aspects of politics.

Unlike previous studies, our research is motivated by the scant evidence on the impact of a divided government, which, in conjunction with election and partisan variables, gauges the relevance of the key theoretical perspectives in the area that these represent i.e., divided government theory (Lohmann & O'Halloran, 1994), partisan business cycle theory, (Nordhaus, 1975; Lobo & Tufte, 1998), and partisan theory (Hibbs, 1994). Nordhaus (1975) explains that governments are driven by opportunistic behaviour[1], while Hibbs (1994) by the rational behaviour of politicians; both explanations are part of the political macroeconomy, yet focus on different driving forces.

Apart from treating the U.S. federal elections as a source of domestic political uncertainty, this study also uses U.S. politics as a proxy for global risk (Brogaard et al., 2020). We qualify the latter by arguing that a) the U.S. elections matter to other countries given their hegemonic position as the world's largest economy and their dominance in global trade; b) the U.S. elections are the most covered globally, and c) partisan politics in the U.S. implies that the outcome generates material uncertainty over policy stemming from unpredictable election outcomes.

Therefore, we propose to revisit the impact of political events and augment its scope by focusing on international politics to understand exchange rate volatility. To this end, by applying a GARCH (1,1) methodology on weekly data spanning the period 2000 to 2021, we explore the impact of U.S. domestic politics on exchange rate volatility of the spot rate of five currencies, i.e. the Japanese yen, the Canadian dollar, the British pound, Mexican peso, and the euro, with the U.S. dollar. We also incorporate three political variable categories based on theories derived from political/international political economy and six macroeconomic control variables. The evidence suggests that the partisan and divided government variables significantly impact the conditional variance equation, whilst the observed reduced levels of exchange rate volatility during a Democrat presidency run counter to prior studies of partisanship (see Lobo & Tufte, 1998). Finally, exchange rate volatility increases one month before an election.

[1] In Nordhaus's (1975) model, governments care about reelection prospects. Therefore, they exploit the Phillips curve, i.e., the trade off between unemployment and inflation, as politicians have little reason to value post-election consumption.

The structure of this paper is as follows: section 2 reviews the literature, whilst section 3 discusses the data and the methodology utilised. Section 4 reports and discusses the empirical results, and section 5 provides some concluding remarks.

2. THEORETICAL PERSPECTIVES

The outcome of U.S. elections can have significant implications for economic policies, international relations, and financial markets. Elections can be viewed as a source of uncertainty or a proxy for risk as the outcome can have a material impact on the global economy given the size and influence of the U.S. economy. One argument favouring this view is that different political parties have different policy preferences, and their control of government institutions can lead to significant changes in the regulatory framework, tax, and spending priorities. For example, the election of Donald Trump in 2016 led to a substantial shift in U.S. trade policy with the imposition of tariffs on imports. Then, the election of Joe Biden in 2020 led to expectations of a more expansionary fiscal policy and higher taxes for corporations and wealthy individuals. Both election results affected global supply chains, prices, and growth expectations.

Several studies argue that U.S. elections affect financial markets and the global economy. For instance, the International Monetary Fund (IMF) regularly assesses the impact of political events, including U.S. elections, on the global economy in its World Economic Outlook (WEO) reports. The World Bank also acknowledges the impact of political events, including U.S. federal elections, on the global economy in its Global Economic Prospects reports.

Similarly, BIS (2020) examines the effect of geopolitical risks on exchange rate volatility in emerging market economies. The authors use a panel data set of 36 emerging market economies in the period 1996–2019 and find that geopolitical risks significantly affect exchange rate volatility.

Generally speaking, scholars have been reluctant to engage vigorously with empirical research exploring political factors' impact on exchange rate volatility. In the extant academic literature, three theoretical strands treat politics as a factor potentially explaining exchange rate volatility. The most impactful, oldest, and widely researched one looks at political events as a source of volatility given the uncertainty concerning future policy (see Bachman, 1992; Lobo & Tufte, 1998).

The second strand engages with political indices as proxies for political risk but often ignores exchange rate volatility (see Melvin & Tan, 1996; dos Santos et al., 2021; Vortelinos & Saha, 2016), whilst the third one looks at the role of socio-political and political instability factors, such as civil unrest, and their impact on exchange rates (see Kutan & Zhou, 1995; Melvin & Tan, 1996; Bouraoui & Hammami, 2017).

2.1 Elections, partisanship, and exchange rates

The most influential and widely researched area of politics and exchange rates looks at political risk by using political events as a proxy, such as elections, partisanship, and approval rating variables. Generally, the literature finds consistent evidence suggesting that elections impact exchange rate volatility (see Lobo & Tufte, 1998; Bachman, 1992; Blomberg & Hess, 1997).

Bachman (1992) by using political news (specifically, elections) as a determinant of bias in the forward market, established that half of the elections were significant for determining forward bias, whilst no discernible pattern as to the influence on exchange rate volatility was detected. Furthermore, Blomberg and Hess (1997), in investigating whether politics influences exchange rates and whether incorporating political variables will improve the accuracy of exchange rate forecasting, found that a) exchange rates are sensitive to political variables regardless of incorporating economic variables; b) there is no evidence to suggest that exchange rates and political variable work in the opposite direction – i.e., exchange rates influence a president's approval ratings; and c) the political variables for partisanship, elections, and approval rating are statistically significant, but, unexpectedly, approval ratings have a negative coefficient, suggesting that a popular president leads to currency depreciation. While this evidence shows that political variables are relevant to determining exchange rates, it does not suggest that this affects exchange rate volatility.

On the impact of political events on exchange rate volatility, Lobo and Tufte (1998) found an effect on conditional variance or volatility for four USD currency pairs (Japanese yen, German mark, British pound, and Canadian dollar) during election years. Although the authors could not discern a clear pattern of either an increase or a decrease in volatility across all pairs, they observed an increase in volatility during a Democrat rather than a Republican government.

Assuming that a) volatility is a consequence of uncertainty relating to government policy mix and b) investors are sensitive to changes in the political regime, then one can infer that in the context of U.S. elections, exchange rate volatility can be linked to the political economy literature via two channels supported by the political business cycle theory (see Nordhaus, 1975) and the partisan theory (see Hibbs, 1994). Nordhaus (1975) explains it in terms of opportunistic behaviour while Hibbs (1994) explains it in terms of rational behaviour of politicians.

According to the political business cycle theory (PBC)[2], politicians are opportunistic and seek to maximise the probability of getting re-elected, which is contingent upon the economy's performance; therefore, governments adopt different policies before elections as opposed to during non-election periods. The partisan theory predicts that politicians will differ in policy outcomes along partisan lines. In turn, the outcome of policy associated with this will be determined by the party's ideological position. These policy decisions are constrained by the economic structure, which is normally represented by the trade-off between inflation and unemployment (the dynamic Philips curve) (Lobo & Tufte, 1998). In this context, Blomberg and Hess (1997) suggest a distinct difference between a left-wing and a right-wing government, as left-wing governments have a higher tolerance of inflation. This is confirmed by Siokis and Kapopoulos (2003), who also found that a left-wing (socialist) party versus an incumbent right-wing party had a significant impact on the conditional variability of the currency (i.e., the Greek drachma at that time).

Additional work by Leblang and Bernhard (2006) found that the conditional variance of the exchange rate depends on the position within the election cycle. Specifically, for both Belgium and Sweden, there was an increase in volatility during the dissolution period, but the post-election period generally saw a decrease in volatility. The authors also argued that the nature of the political system needs to be considered when designing a model to explore political variables.

[2] It should be noted that PBC leads to a great extent to the explanation of political (electoral) cycles that leads on to political and economic instability and then possibly to exchange rate volatility as a consequence either of opportunistic policy or partisan policy.

Further evidence suggests that countries with majoritarian systems experience lower levels of volatility than pluralist systems at various stages in the electoral cycle. For instance, France and Britain (majoritarian systems) saw no consistent volatility during the dissolution of governments, whilst Sweden and Belgium (pluralist systems) saw increases during negotiations and dissolution periods. This is due to the higher uncertainty of a strong government being formed under a pluralist system, as coalitions are more likely, leaving room for policy uncertainty.

In the U.S., however, electoral periods are exogenous as they are set out in the constitution and are therefore not necessarily subject to the same constraints, such as the predictability of the parliamentary system. In contrast, the U.K. elections are found to often be accompanied by periods of lower volatility due to the predictability of the British parliamentary system (Leblang & Bernhard, 2006). Furthermore, Garfinkel et al. (1999) established that surprising or unforeseen election outcomes significantly raised uncertainty for investors, whilst anticipated or ambiguous election results were insignificant. Freeman (2000) provides evidence that links democratic systems to electoral outcomes and finds significance in determining exchange rate volatility due to increased uncertainty over policy outcomes. In the same spirit, the results of Siokis and Kapopoulos (2003) reinforce the view established in prior research on the importance of partisanship and elections in affecting currency fluctuations.

Leblang and Bernhard (2006), whilst not directly researching the role of hegemonic politics, do include foreign politics as a variable. More specifically, their study on the role of German elections in exchange rate volatility for four European currency pairs found inconsistent evidence to suggest that German elections influence foreign currency pairs. The German polls did not have any significant impact on the British pound. For the Belgian and French franc, however, the campaign period proved to be significant, exhibiting a negative coefficient, i.e., signalling lower volatility during elections, whereas the German period of negotiation had a positive and significant impact on the volatility of Swedish krona.

Liu and Pauwels (2012) approach the study of international politics and exchange rates by examining whether international political pressure influences the

Chinese exchange rate (CNY/USD) through an events study methodology. They divide political pressure into U.S. political pressure and non-US aggregated pressure[3]. Their findings suggest that the aggregate external pressure and U.S. pressure indicators positively affect exchange rate volatility. Another significant factor is the negative coefficient of the Sino-US meetings indicator, suggesting that volatility declines during diplomatic talks between the two countries. The distinction between these coefficients is that the market reacts to uncertainty surrounding policy between two nations and prefers dialogue. Furthermore, Brogaard et al. (2020) investigated the role of global political uncertainty across asset classes, primarily domestic stock markets but also sovereign bonds and foreign exchange.

Following the rationale of the prior literature, this study treats U.S. federal elections as a source of domestic political uncertainty and as a proxy for global risk. Overall, evidence suggests that U.S. elections generally lead to an appreciation of the U.S. dollar against foreign currencies, whilst insights into exchange rate behaviour are limited (Menkhoff et al., 2012, 2017).

2.2 Political risk indices

The next strand of the literature investigates political risk proxied by compound indices. Using political risk indices to analyse foreign exchange has become popular in recent years. This is potentially due to the extensive focus of earlier research on connecting political event variables with currency volatility owing to the difficulties associated with using political risk indices. Fundamentally, most of these variables are very subjective, making it challenging to consistently and accurately apply numerical values to issues such as judicial independence. This can be seen as a reason for studying events; subjectiveness is not necessarily a major concern with event studies due to the binary nature of using events such as elections.

Despite these challenges, however, indices have been developed and effectively used in the respective literature. Melvin and Tan (1996), dos Santos et al. (2021),

[3] The definition of political pressure given by Liu and Pauwels (2012) stems from public statements from officials on Chinese Exchange Rate Policy from the US, EU, Japan, and major international organisations, such as the International Monetary Fund, the G7 Group, and the Asian Development Bank.

Vortelinos and Saha (2016) and Bekaert et al. (2014), using the International Country Risk Guide index (ICRG) and the Economic Policy Uncertainty index (EPU), concluded that political risk is significant and positively correlated with exchange rate volatility and returns. More specifically, Melvin and Tan (1996), using the ICRG index in a cross-sectional setting, found a connection between higher political risk and a higher spread for the South African rand, which reflects higher market uncertainty and volatility. Similarly, Vortelinos and Saha (2016) found that political risk had a significant impact on the volatility of exchange rate markets, particularly in North America, the most affected.

Furthermore, Filippou et al. (2017) and Brogaard et al. (2020), when exploring how political shocks influence currency investment strategies and their profitability, found that unexpected global political risk is priced into the cross-section of currency momentum strategies and explains a lot of the excess returns. Notably, the authors conclude that speculators will demand a premium on currencies with significant exposure to global political risk but specifically U.S. political shocks, especially in the short term. This generally highlights that there is a foundation in the literature for U.S. risk or variables being used as a proxy for global risk.

In a nutshell, the extant literature has provided sound evidence demonstrating that compounded political indices can be used to study political risk as an alternative to event studies of politics and exchange rates.

2.4 Socio-political instability.

The third and final strand of the literature examines socio-political factors as explanatory variables for exchange rate volatility. Several studies have established that socio-political events are linked to higher exchange rate volatility (see Kutan & Zhou, 1995; Melvin & Tan, 1996; Bouraoui & Hammami, 2017). In particular, Kutan and Zhou (1995), linking socio-political factors to higher exchange rate volatility in post-Communist Poland, find that socio-political unrest increases exchange rate volatility. Country risk (a proxy for political unrest) is captured by a dummy variable based on news stories relating to events reflecting poor socio-political conditions and social unrest, such as strikes, violent demonstrations, political talks, and elections.

In a different study, Melvin and Tan (1996), by exploring the impact of socio-political unrest resulting from racial policies on the volatility of the South African rand, found that its conditional variance was significantly affected by the envisaged socio-political factors. Similarly, Proti (2013) found a significant but negative relationship by investigating exchange rate fluctuations and political tensions in Tanzania. This was due to fear of a new regime's uncertain economic policy, which might affect investors' funds. Similar evidence was established by Saeed et al. (2012) in a study of Pakistan, where political instability was a significant factor that caused the depreciation in Pakistan's exchange rate.

Additional evidence on the relationship between political instability (social unrest) and currency fluctuations are provided by Bouraoui and Hammami (2017), who focused on five countries involved in the Arab Spring. By creating a political instability index to incorporate the frequency of government change, terrorist attacks, revolutionary disorder, regional conflict, and assassinations split into long- and short-term impacts, Bouraoui and Hammami (2017) found that political instability was a significant determinant of exchange rate fluctuations for Egypt and Tunisia in the short term.

There seems to be a consensus that an increase in political instability or socio-political problems will likely increase exchange rate volatility in emerging markets. Even though the U.S. does not necessarily exhibit the same fundamental instability as these developing counties, we can draw several interesting conclusions from the literature that will help in shaping the scope of this study.

Overall, the foundational literature tends to focus on political events, with the use of indices becoming a more viable alternative today. There is also evidence to support the impact of domestic political variables on exchange rates, mainly focused on elections. However, the fact that there has been some notable disregard for partisan variables in the literature, in conjunction with the limited focus on global politics, highlights a genuine scope for further investigation into the link between international politics and exchange rate volatility.

3. DATA AND METHODOLOGY

The data set contains 1096 observations spanning ten federal elections (five presidential and five midterms), covering the first week of January 2000 to the first week of January 2021, and the spot rates for five currency pairs: JPY/USD, GBP/USD CAD/USD, PESO/USD. EURO/USD[4]. These are weekly Wednesday-close data in the spirit of Siokis and Kapopoulos (2003) sourced from the Refinitiv Eikon Datastream. By using three currency pairs with close but varying levels of economic interdependence with the United States, we aim to explore the linkages between international relations/international political economy and exchange rate volatility.

It is well known that the U.S. has many trading partners worldwide, but some countries are more significant trade partners than others. For instance, Canada and Mexico are the top two trading partners of the USA due to their proximity and membership of the North American Free Trade Agreement (NAFTA). Notably, the United States makes up 75.37% of Canadian exports and imports, making it Canada's largest trading partner by a significant margin (WITS, 2021a). Japan, Germany, and France are also major trading partners, especially in the automotive and technology industries, which explains the inclusion of the yen and the euro in the currency pairs. The U.S is the second-largest trade partner for Japan, after China, accounting for 19% of trade (WITS, 2021c). Finally, the U.K. is a key trading partner of the U.S., with a significant volume of goods and services traded between the two countries; the U.S. is the largest exporter market for the U.K., accounting for 15.07% of trade (WITS, 2021b). Overall, the U.S. and the U.K. have a long-standing historical relationship that has shaped the political and economic ties between the two countries and continues to impact their relationship.

In Table 1, we report the summary statistics for the weekly return while the preliminary data analysis indicates the presence of a unit root in the series[5].

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4 In this study, we focus on the top five trading partners of the US, hence the corresponding selection of the five currencies.

5 For economy of space, we do not report the ADF tests, but they are available upon request.

Table 1. Summary statistics

	JPY return	CAD return	GBP return	PESO return	EURO return
Obs	1,096	1,096	1,096	1,096	1,096
Mean	0.0015	0.0113	0.0163	0.000156	0.000019
StDev.	1.3869	1.2061	1.3665	0.0071	0.0058

In our empirical analysis, we use the volatility of exchange rate returns as the dependent variable. In line with Brogaard et al. (2021), who use U.S. domestic politics as a proxy for global risk in their study on the determinants of exchange rates, we utilise similar methods by delving deeper into the literature on international political economy and bringing into play several new 'political' variables considered to be important. To explore the role of U.S. domestic politics as a proxy for global influence or their role in the international environment, we use political events as proxies for domestic political change or risk.

The first set of political variables are the three electoral variables designed in the spirit of Brogaard et al. (2021), similar to the relevant literature (Lobo & Tufte, 1998; Leblang & Bernhard, 2006) and covering ten federal election cycles – five presidential and five midterm elections. The three electoral variables reflect a six-month period prior to the election, a three-month period prior to the election, and a one-month period prior to the election. Earlier literature (Lobo & Tufte,1998; Siokis & Kapopoulos, 2003) uses only one electoral variable, which is set 16 weeks before the election period, whereas Leblang and Bernhard (2006) take a more nuanced approach and argue for the significance of differentiating between the periods of the electoral cycle. In the same spirit as Leblang and Bernhard (2006), we distinguish between periods before the electoral cycle. In detail, six months prior to the election is effectively the start of the election cycle with six months covering most of the primaries, when the outcomes of primaries become clear and the candidates are essentially decided. Three months is consistent with the time when the Democrat and Republican conventions meet to formally decide who the candidates are. However, as pointed out by Brogaard et al. (2020), this is a relatively quiet period. One month prior to the election is the final month of campaigning alongside the presidential debates and the eventual vote, arguably the most significant period before an election[6].

The next political variable used is the partisan variable, which follows the same pattern as Lobo and Tufte (1998) based on the rationale of the partisan cycle (Hibbs, 1994). The dummy variable assumes the value of 1during a Democrat presidency and 0 during a Republican presidency[7].

As established in the work of Lohmann and O'Halloran (1994), another potential factor influencing domestic politics or U.S. foreign policy is divided government. This is where a party different to that of the president controls the Congress. As such, we create a divided government dummy variable coded as 1 during periods of divided government and 0 otherwise. As there are two houses of Congress, there can be periods, i.e. the 2001–2003 Congress, where the House of Representatives was Republican, the president was Republican, but the Senate was divided 50/50. In this case, since the vice president, as president of the Senate, has the deciding vote in the Senate, the Senate was classified as being Republican, so the government was classified as not being divided. We expect higher levels of volatility during periods of divided government. The data for U.S. elections and U.S. government composition are taken from the Federal Election Commission and Brookings, respectively.

In line with prior literature (Liu & Pauwels, 2012; Neely, 2005), we also control for the macroeconomic environment. Whilst the literature does not entirely agree on what causes exchange rate volatility, there is some consensus on the critical factors. The first macroeconomic factor included in the model is the OECD recession indicator for Japan, Canada, Mexico, the eurozone, and the U.K. This factor controls for the economic cycle, economic downturns, and recessions. Similarly, we include shocks to the monetary system based on the money supply proxied by M2 (Liu & Pauwels, 2012) but also add the consumer price index (CPI). The money supply (M2) controls for economic and monetary conditions,

6 A fact becoming even more embedded by the 84.4 million people who tuned in for the Trump–Clinton debate, the largest viewership in US history (Statista, 2020).

7 Lobo and Tufte (1998) found that periods of Democrat presidency were met with higher periods of volatility so, in line with their previous findings, we have assigned the value of 1 to a Democrat presidency.

but also, as Bouraoui and Hammami (2017) point out, stands as a transmission mechanism for political events. Monetary growth shocks have also been reported as having an influence on exchange rate volatility in previous research (Mpofu, 2016; Bouraoui & Hammami, 2017; Junttila & Korhonen, 2012). The fourth factor incorporated in the model is the trade environment proxied using trade openness, a commonly used factor in the literature derived from Hau (2002) and utilised by Stancik (2007). Hau (2002) finds a negative and significant relationship between volatility and trade openness, suggesting that higher levels of trade integration lead to a lower level of exchange rate volatility, broadly in line with Mpofu (2016), Stancik (2007), and Suleman and Berka (2017).

Trade openness is also an important variable for our political analysis as hegemonic stability theory finds a connection between trade openness and economic integration. This also helps to incorporate Lohmann and O'Halloran's theory (1994) that divided government leads to lower trade openness through a more protectionist trade policy, which results in potentially higher exchange rate volatility. The fifth control variable used to explain the exchange rate volatility is the volatility of oil prices. Specifically, the oil price enters the equation through the log of WTI Crude as one of the core commodity benchmarks. This is viewed as an important consideration for countries such as Canada and the U.K. Canada represents the world's third largest exporter of oil, accounting for 8% of global oil exports, with 98% going to the United States . Equally, oil is the U.K.'s fifth largest export (IEA, 2021). As Mpofu (2016) points out, commodities can be used as an accurate measurement of terms of trade given the availability of higher frequency data. Thus, the oil price adds robustness to our understanding of the effects of trade conditions in both Canada and the U.K[8]. In line with prior literature (Mpofu, 2016; Siokis & Kapopoulos, 2003), we also control for the non-stationarity in the data by transforming the time series accordingly.

The next step involves establishing the specification of the GARCH(1,1) model. Typically, the literature on exchange rate volatility uses ARCH or GARCH models in dealing with conditional variance (Neeley, 2005). The conditional

8 See Table A in the Appendix for the definition of variables and sources. Where applicable, we used linear conversion to obtain the same frequency in the dataset.

variance models, as opposed to standard deviation, are superior at dealing with uncertainty, which is relevant here given the unpredictable element of volatility.

The general specification of the model, which is in line with Neeley, (2005), Liu and Pauwels, (2012), and Mpofu (2016), is outlined below:

$$r_t = \gamma_0 + \gamma_1 \Delta r_{t-1} + \varepsilon_t \tag{2}$$

Whilst higher-order specifications do exist, the present specification is generally accepted to be effective and parsimonious (Mpofu, 2016). The GARCH term refers to the persistence of the volatility or the long-run volatility. Strictly speaking, providing $\alpha_0 \geq 0$, $\alpha_1 \geq 0$, $and\ \beta_1 \geq 0$, we will have positive conditional variance. Furthermore, by providing $\alpha_1 + \beta_1 < 1$, there is sufficient evidence for a second moment in the equation (Liu & Pauwels, 2012). The final specification for model outlined in equation (2) is as follows:

$$r_t = \gamma_0 + \gamma_1 \Delta_{r-1} + \zeta_1 USelec1_t + \zeta_2 USelec3_t + \zeta_3 USelec6_t + \zeta_4 Divgov_t + \zeta_5 Partisan_t + \zeta_6 Reces_t + \zeta_7 M2_t + \zeta_8 CPI_t + \zeta_9 Oil_t + \zeta_{10} UST\text{ROP}_t + \zeta_{11} T\text{ROP}_t + \varepsilon_t \tag{3}$$

$$\text{Var}(\varepsilon_t) = \sigma_t^2 = a_o + a_1 \varepsilon_{t-1}^2 + \beta_1 \sigma_{t-1}^2 + \text{Exp}(\zeta_1 \text{USelec1}_t + \zeta_2 \text{USelec3}_t + \zeta_4 \text{Divgov}_t + \zeta_5 Partisan_t + \zeta_6 Reces_t + \zeta_7 M2_t + \zeta_8 CPI_t + \zeta_9 Oil_t + \zeta_{10} UST\text{ROP}_t + \zeta_{11} T\text{ROP}_t), \tag{4}$$

where r_t denotes the volatility of exchange rate returns; Δ_{r-1} is the first difference of the lagged value of exchange rate volatility; $USelec1_t$, $USelec3_t$, and $USelec6_t$ denote the three electoral variables reflecting a one-month period, a three-month period, and a six-month period prior to the election; $Partisan_t$ is the partisan variable that takes the value of 1during a Democrat presidency and 0 during a Republican presidency; $Divgov_t$, is the divided government dummy, which assumes the value of 1 during periods of divided government and 0 otherwise; $Reces_t$ is OECD recession indicator; $M2_t$ is the money supply; CPI_t is a measure of inflation; Oil_t is the volatility of oil prices; $UST\text{ROP}_t$ denotes U.S. trade openness whereas $T\text{ROP}_t$ denotes trade openness of the respective economies (Japan, the U.K., and Canada) which alternate in the three different estimated models. Equation 4 is the standard specification of the variance equation in the GARCH(1,1) specification.

In line with past literature (see Lobo & Tufte, 1998; Siokis & Kapopoulos, 2003; Liu & Pauwels, 2012), we include the independent variables in the mean return equation. To ensure robustness, we repeat the regression individually, removing and adding the election dummy variables to account for collinearity moving from the one-month, then adding the three-month, and finally looking at the six-month alone (see Tables 2, 3 and 4).

4. EMPIRICAL RESULTS

We report the results of the regressions in Tables 2, 3, and 4 for the currency pairs CAD/USD, GBP/ USD, and JPY/USD, respectively. We determine the best-fit model based on the AIC and BIC criteria, and we use robust standard errors to improve the robustness of the results. In line with the literature, our results on the implications of politics or international political economy on exchange rate volatility are mixed. While it seems that the political variables have little to no effect on the mean exchange rate, evidence suggests that these variables are significant for the conditional variance. At this point, we note that our paper aims to reveal the impact of political variables on exchange rate volatility and not the determinants of the exchange rate by focusing on the conditional variance equation.

Among the political variables, the election variables present the most indistinct results. Generally, one should expect countries with the highest level of economic integration to be the most impacted by exchange rate volatility during an election season. However, our results suggest this was not the case in the examined period. Starting with Canada in Table 2, the election dummy variables for the first, third, and sixth months were all insignificant for the mean return equation. The result is consistent with previous studies (see, for example, Siokis & Kapopoulos, 2003; Lobo & Tufte, 1998; Leblang & Bernhard, 2006) where electoral variables were found to be insignificant in the mean equation. However, the conditional variance tells a different story.

While the six-month and three-month variables were both insignificant, the one-month variable was significant at the 10% level, which is in line with Lobo and Tufte (1998). This evidence is also consistent with both Leblang and Bernhard (2006) and Siokis and Kapopoulos (2003), who investigated different pairs of currencies. The large change in the coefficient size from models 1 to 2 suggests that the scale of this increase in volatility is unclear and inconsistent. However, for Canada, the signs of the coefficients are interesting in that we see a positive coefficient for one month, suggesting an increase in volatility, a negative coefficient for three months, and a positive coefficient for six months. Although the six and three months were insignificant, the general pattern is consistent with what was expected. During the primaries and presidential debates, we see more uncertainty and higher volatility, but during a quiet period three months prior, the volatility is lower.

Table 2: CAD/ USD regression results

Model	1		2		3		4	
	Coefficient	St. Error	Coefficient	St. Error	Coefficient	St. Error	Coefficient	St. Error
Mean Equation								
Intercept	-1.865	2.071	-2.646	2.104	-2.608	2.109	-2.534	2.113
Political Variables								
US election 1 month	0.133	0.179	0.075	0.145				
US election 3 month	-0.031	0.147			0.013	0.105		
US election 6 month	-0.024	0.090					-0.015	0.069
Divided Government	-0.007	0.103	0.020	0.107	-0.007	0.089	0.012	0.106
Partisan	0.147*	0.855	0.135*	0.094	0.1457*	0.093	0.145*	0.094
Macroeconomic Controls								
Recession	-0.050	0.071	-0.039	0.074	-0.032	0.073	-0.032	0.073
US Trade Openess	0.509	0.529	0.266	0.561	0.300	0.559	0.312	0.557
CAN Trade Openess	1.431**	0.682	1.342**	0.684	1.369**	0.681	1.363**	0.682
Inflation	0.607	0.439	0.691*	0.424	0.696	0.427	0.684*	0.427
Crude Oil	-3.989***	0.493	-4.601**	0.648	-4.590***	0.636	-4.592***	0.634
Money Supply (M2)	2.118***	1.481	2.156***	1.539	2.149***	1.540	2.154***	1.544
Conditional Variance								
Intercept	13.923*	7.520	13.060*	7.480	13.518*	7.343	13.809*	7.579
Political Variables								
US election 1 month	1.817*	1.316	0.376*	0.323				
US election 3 month	-1.340	1.335			-0.184	0.315		
US election 6 month	0.042	0.397					-0.127	0.253
Divided Government	0.597*	0.375	0.652*	0.376	0.545*	0.361	0.557	0.368
Partisan	-0.559**	0.298	-0.603**	0.297	-0.497**	0.283	-0.509	0.284
Macroeconomic Controls								
Recession	0.170	0.227	0.159	0.228	0.207	0.222	0.207	0.225
US Trade Openess	-0.364	2.134	-0.425	2.101	-0.404	2.078	-0.313	2.118
CAN Trade Openess	-5.279**	2.376	-5.088**	2.422	-5.219**	2.302	-5.139*	2.323
Inflation	-4.063***	1.447	-3.891***	1.479	-3.968***	1.431	-4.002***	1.469
Crude Oil	-9.872***	2.821	-10.347***	2.763	-9.961***	2.749	-9.920***	2.779
Money Supply (M2)	2.278	1.979	1.772	0.332	2.157	1.633	2.174	1.656
ARCH & GARCH								
ARCH	0.100***	0.0249	0.102***	0.0241	0.106***	0.023	0.106***	0.023
GARCH	0.820***	0.3562	0.821***	0.0337	0.812***	0.035	0.814***	0.034
AIC	3225.958		3219.460		3220.908		3221.105	
BIC	3355.943		3329.450		3330.895		3331.096	
PseudoLoglikelihood	-1586.227		-1587.730		-1588.454		-1588.553	
Number of observations	1096		1096		1096		1096	
WaldChi2(11)(9)	88.66		89.33		89.42		89.42	

Notes: ***, **, * denote significance at the 1%, 5%, and 10% significance levels, respectively.

Table 3: GBP/ USD regression results

Model	1		2		3		4	
	Coefficient	St. Error	Coefficient	St. Error	Coefficient	St. Error	Coefficient	St. Error
Mean Equation								
Intercept	-1.561**	0.726	-1.446**	0.729	-1.484**	0.721	-1.577**	0.722
Political Variables								
US election 1 month	-0.050	0.217	0.095	0.186	0.585	0.241		
US election 3 month	0.087	0.153			0.133	0.106		
US election 6 month	0.082	0.109					0.114*	0.082
Divided Government	0.360***	0.119	0.335***	0.118	0.351***	0.117	0.357***	0.117
Partisan	0.018	0.104	0.035	0.103	0.027	0.103	0.019	0.104
Macroeconomic Controls								
Recession	0.061	0.091	0.071	0.088	0.073	0.088	0.056	0.090
US Trade Openess	-1.618**	0.704	-1.487**	0.706	-1.514**	0.700	-1.645**	0.701
UK Trade Openess	1.618**	0.712	1.482**	0.713	1.501**	0.710	1.651**	0.710
Inflation	2.046	1.912	2.050	1.908	2.045	1.911	2.047	1.912
Crude Oil	-2.942***	0.613	-2.952***	0.616	-2.938***	0.613	-2.947***	0.611
Money Supply (M2)	1.817***	1.379	1.830***	1.372	1.821***	1.373	1.820***	1.379
Conditional Variance								
Intercept	-5.737***	1.539	-5.502***	1.505	-5.605***	1.507	-5.779	1.502
Political Variables								
US election 1 month	0.112*	0.085	0.164*	0.036				
US election 3 month	-0.023	0.306			-0.030	0.301		
US election 6 month	0.220	0.210					0.186*	0.138
Divided Government	0.085	0.204	0.072	0.200	0.074	0.202	0.090	0.198
Partisan	0.398**	0.170	0.428***	0.162	0.424***	0.155	0.399***	0.149
Macroeconomic Controls								
Recession	0.378**	0.171	0.421***	0.163	0.404	0.166	0.381**	0.164
US Trade Openess	-2.703***	1.449	-2.465***	1.422	-2.598	1.420	-2.729**	1.409
UK Trade Openess	0.078	1.317	-0.145	1.319	0.002	1.327	0.061	1.297
Inflation	2.987***	1.309	2.967***	1.298	2.966***	1.307	2.970***	1.303
Crude Oil	-11.231***	1.689	-11.043***	1.664	-11.002***	1.357	-11.126	1.271
Money Supply (M2)	1.875	1.807	1.880	1.865	1.863	1.845	1.866	1.790
ARCH & GARCH								
ARCH	0.082***	0.028	0.087***	0.025	0.091***	0.026	0.0831***	0.025
GARCH	0.706***	0.057	0.696***	0.054	0.695***	0.052	0.705***	0.049
AIC	3571.807		3567.653		3569.924		3564.221	
BIC	3701.792		3677.64		3689.910		3674.208	
PseudoLoglikelihood	-1759.9		-1761.827		-1760.962		-1760.110	
Number of observations	1096		1096		1096		1096	
WaldChi2(11)(9)	39.71		38.29		41.00		40.33	

Notes: ***, **, * denote significance at the 1%, 5%, and 10% significance levels, respectively.

Table 4: JPY/USD regression results

Model	1		2		3		4	
	Coefficient	St. Error	Coefficient	St. Error	Coefficient	St. Error	Coefficient	St. Error
Mean Equation								
Intercept	-0.025	0.800	0.106	0.787	0.108	0.786	0.098	0.799
Political Variables								
US election 1 month	-0.188	0.213	-0.081	0.195				
US election 3 month	0.273*	0.154			0.074	0.130		
US election 6 month	-0.145	0.109					-0.065	0.092
Divided Government	0.090	0.107	0.044	0.108	0.044	0.109	0.064	0.106
Partisan	0.060	0.122	0.052	0.113	0.047	0.114	0.079	0.115
Macroeconomic Controls								
Recession	-0.062	0.082	-0.056	0.082	-0.049	0.081	-0.071	0.081
US Trade Openess	-0.069	0.758	0.067	0.747	0.067	0.748	0.012	0.752
JPN Trade Openess	0.090	0.311	0.051	0.317	0.066	0.318	0.099	0.319
Inflation	-3.495	71.920	5.601	69.153	3.109	68.943	-3.383	69.237
Crude Oil	-0.191	0.811	-0.062	0.828	-0.047	0.826	-0.070	0.840
Money Supply (M2)	1.906**	0.811	1.908**	0.828	1.930***	0.826	1.887**	0.840
Conditional Variance								
Intercept	-7.173***	2.024	-3.796***	1.502	-3.599*	1.491	-3.392**	1.442
Political Variables								
US election 1 month	1.560*	0.857	0.693***	0.248				
US election 3 month	1.319***	0.392			0.125	0.181		
US election 6 month	-1.381***	0.719					-0.221*	0.149
Divided Government	1.346***	0.330	0.681***	0.205	0.730***	0.212	0.739***	0.204
Partisan	-0.473**	0.243	-0.247*	0.163	-0.273*	0.163	-0.237*	0.156
Macroeconomic Controls								
Recession	0.000	0.214	-0.233	0.158	-0.250*	0.157	-0.266*	0.156
US Trade Openess	-3.529**	1.647	-1.282	1.220	-1.132	1.193	-1.030	1.155
JPN Trade Openess	0.180	0.680	-0.198	0.522	-0.232	0.510	-0.239	0.508
Inflation	-5.313	4.053	-3.793	2.584	-4.081	2.527	-4.934*	2.575
Crude Oil	-2.166	4.163	-4.278*	2.589	-4.967**	2.352	-5.130	2.207
Money Supply (M2)	1.992	1.188	1.990	1.278	1.995	1.260	1.847	1.268
ARCH & GARCH								
ARCH	0.070***	0.017	0.086***	0.022	0.088***	0.022	0.085***	0.022
GARCH	0.878***	0.021	0.804***	0.039	0.801***	0.040	0.802***	0.039
AIC	3744.698		3754.397		3758.942		3757.469	
BIC	3869.683		3864.386		3868.930		3867.457	
PseudoLoglikelihood	-1847.349		-1857.471		-1857.471		-1856.735	
Number of observations	1096		1096		1096		1096	
WaldChi2(11)(9)	11.6		9.96		11.09		11.39	

Notes: ***, **, * denote significance at the 1%, 5%, and 10% significance levels, respectively.

Table 5: PESO/USD regression results

Model	1		2		3		4	
	Coefficient	St. Error	Coefficient	St. Error	Coefficient	St. Error	Coefficient	St. Error
Mean Equation								
Intercept	-2.673	3.134	-2.893	3.732	-3.085	3.984	-3.794	3.749
Political Variables								
US election 1 month	0.023	0.297	0.066	0.158				
US election 3 month	0.084	0.313			0.027	0.165		
US election 6 month	-0.062	0.193					-0.183	0.167
Divided Government	-0.001	0.203	0.031	0.207	-0.019	0.185	0.175	0.216
Partisan	0.062***	0.005	0.216***	0.019	0.192***	0.018	0.092***	0.001
Macroeconomic Controls								
Recession	-0.021	0.172	-0.163	0.963	-0.096	0.098	0.099	0.172
US Trade Openess	0.197	0.324	0.162	0.199	0.168	0.322	0.172	0.287
MEX Trade Openess	0.789***	0.083	1.253*	0.774	1.026**	0.472	1.037**	0.494
Inflation	0.028	0.672	0.189	0.924	0.563	0.893	0.427	0.703
Crude Oil	-3.681**	0.034	-3.389***	0.063	-3.680***	0.037	-3.246***	0.08
Money Supply (M2)	35.993***	15.784	22.794***	4.603	19.784***	14.392	16.678***	9.683
Conditional Variance								
Intercept	8.784***	3.393	8.089***	2.643	8.584***	3.684	8.694*	2.683
Political Variables								
US election 1 month	0.589*	0.325	0.457***	0.199				
US election 3 month	-0.983	0.998			-0.563	0.468		
US election 6 month	0.183	0.828					-0.093	0.159
Divided Government	0.087***	0.015	0.052***	0.003	0.035***	0.011	0.673	0.583
Partisan	-0.362***	0.019	-0.273***	0.117	-0.268**	0.028	-0.119	0.343
Macroeconomic Controls								
Recession	0.189	0.364	0.194	0.246	0.327	0.294	0.301	0.273
US Trade Openess	-0.274	1.484	-0.392	1.633	-0.203	1.123	-0.124	1.263
MEX Trade Openess	-1.889**	0.034	-1.374***	0.384	-1.673***	0.232	-1.282***	0.182
Inflation	-2.164***	0.183	-2.744***	0.639	-2.733***	0.431	-2.116***	0.372
Crude Oil	-3.263***	0.784	-4.642***	1.012	-4.477***	1.153	-4.573***	1.112
Money Supply (M2)	13.366	10.494	12.433	8.673	12.933	10.573	11.477	9.673
ARCH & GARCH								
ARCH	0.113***	0.003	0.167***	0.009	0.128***	0.019	0.166***	0.013
GARCH	0.639***	0.162	0.621***	0.027	0.673***	0.051	0.676***	0.018
AIC	3739.673		3763.782		3730.892		3739.173	
BIC	3248.263		3294.378		3274.538		3291.783	
PseudoLoglikelihood	-1467.638		-1468.733		-1469.462		-1468.738	
Number of observations	1096		1096		1096		1096	
WaldChi2(11)(9)	78.47		78.83		78.45		78.49	

Notes: ***, **, * denote significance at the 1%, 5%, and 10% significance levels, respectively.

Table 6: EURO/USD regression results

Model	1		2		3		4	
	Coefficient	St. Error	Coefficient	St. Error	Coefficient	St. Error	Coefficient	St. Error
Mean Equation								
Intercept	-3.783**	1.253	-3.673***	0.247	-3.356***	0.326	-3.382**	0.189
Political Variables								
US election 1 month	-0.023	0.132	0.053	0.234	0.273	0.263		
US election 3 month	0.025	0.231			0.184	0.194		
US election 6 month	0.462	0.305					0.236***	0.073
Divided Government	0.157***	0.0147	0.187***	0.083	0.189***	0.068	0.178***	0.037
Partisan	0.194	0.203	0.283	0.193	0.274	0.178	0.264	0.874
Macroeconomic Controls								
Recession	0.034	0.194	0.036	0.174	0.189	0.199	0.174	0.178
US Trade Openess	-0.461***	0.045	-0.774***	0.046	-0.639***	0.017	-0.584***	0.015
EU Trade Openess	0.873***	0.112	0.872**	0.143	0.419***	0.01	0.138***	0.09
Inflation	5.844	4.983	6.148	6.84	8.484	8.894	11.734	8.844
Crude Oil	-1.592***	0.016	-1.458***	0.014	-1.874***	0.015	-1.788***	0.021
Money Supply (M2)	45.694	33.984	43.834	33.894	46.844	33.743	36.783	28.784
Conditional Variance								
Intercept	-3.783***	1.384	-2.893***	1.207	-3.783***	1.487	-3.373***	1.082
Political Variables								
US election 1 month	0.153***	0.073	0.153***	0.046				
US election 3 month	-0.033	0.366			-0.047	0.561		
US election 6 month	0.38	0.516					0.176*	0.128
Divided Government	0.095	0.304	0.066	0.256	0.089	0.258	0.078	0.204
Partisan	0.467***	0.109	0.578***	0.247	0.498***	0.135	0.482***	0.123
Macroeconomic Controls								
Recession	0.057**	0.007	0.012***	0.003	0.041***	0.006	0.035**	0.004
US Trade Openess	-1.387***	0.491	-1.545***	0.282	-1.874***	0.48	-1.279***	0.462
EU Trade Openess	0.278	1.747	-0.193	1.983	0.004	1.734	0.073	1.279
Inflation	7.441***	2.483	7.793***	2.483	7.083***	2.983	7.997***	2.744
Crude Oil	-3.376***	0.783	-3.238***	0.453	-3.277***	0.373	-3.473	0.981
Money Supply (M2)	24.893	17.833	25.83	14.893	26.916	20.833	21.003	19.8093
ARCH & GARCH								
ARCH	0.063***	0.018	0.099***	0.013	0.088***	0.011	0.078***	0.012
GARCH	0.678***	0.023	0.786***	0.038	0.783***	0.013	0.892***	0.029
AIC	3743.807		3589.399		3673.893		3788.893	
BIC	3701.792		3677.474		3689.973		3674.373	
PseudoLoglikelihood	-1698.993		-1698.363		-1698.673		-1698.673	
Number of observations	1096		1096		1096		1096	
WaldChi2(11)(9)	26.45		28.47		30.99		30.47	

Notes: ***, **, * denote significance at the 1%, 5%, and 10% significance levels, respectively.

The results for the U.K. shown in Table 3 are similar to those of Canada. As expected, none of the electoral dummy variables for the mean equation were significant. Like Canada, the U.K. demonstrates that the one-month U.K. variable was significant at the 10% level in the conditional variance equation, which is in line with Leblang and Bernhard (2006). However, we also find that the coefficients change significantly between models 1 and 2 for the U.K. Beyond this, the 3-month and 6-month variables were both insignificant and remained insignificant in model 3 and 4. The fact that there was evidence of significance for U.S. elections on GBP/USD and higher resulting volatility is generally consistent with Lobo and Tufte (1998). Moreover, we find that all three variables followed the same pattern of coefficient signs as in Canada.

Japan presents the most puzzling results compared to Canada and the U.K., as shown in Table 4. In the mean return equation, there was a significant increase in returns for JPY/USD for the three-month variable, which suggests higher returns in this period. However, this is only in the case of model 1, since it turned out to be insignificant in model 3. The importance of this variable is inconsistent with prior research (Lobo & Tufte, 1998). The conditional variance model provides more certain results than Canada and the U.K. In model 1, all the electoral variables were significant, with the six-month and three-month at the 1% level and the one-month at the 10% level. In model 2, the one-month was significant at the 1% level. Similarly, in model 4, the six-month variable was significant at the 10% level whilst in model 3, the three-month variable was insignificant. Ultimately, there is sufficient evidence to claim that for Japan, the one-month and six-month election variables were significant. Like Canada and the U.K., the one-month was positive, so we see higher volatility before an election. This finding is again consistent with Lobo and Tufte (1998). The significant six-month variable displays a negative coefficient, a surprising outcome considering our expectation of higher volatility six months prior.

In the case of Mexico and the eurozone, Tables 5 and 6, the election dummies for the one-, three-, and six-month variables were all insignificant in the mean return equation, as was the case for Canada. In the conditional variance equation, however, the one-month variable was significant at the 10% and 1% levels, respectively, in line with Lobo andTufte (1998).

Generally, an important pattern can be discerned in that the five currency pairs displayed positivesignificant coefficients in the conditional variance of our one-month election variable. This suggests an increase in volatility one month before a U.S. election, a finding consistent with prior findings. Additionally, it is relatively clear that proximity to elections was not significant in our mean return equation. However, in the conditional variance equation, no discernible pattern was established for the three- and six-month variables. Furthermore, we found little evidence to suggest an increase in the impact of U.S. elections based on economic interdependence, as Japan generally displayed the highest significance levels across our variables.

The partisan variable provided relatively consistent and intriguing results. Based on previous studies (Siokis & Kapopoulos, 2003; Lobo & Tufte, 1998), we expected the partisanvariable to be insignificant in the mean return equation but significant and positive in the conditional variance equation. Across all five currency pairs, the partisan variable was insignificant in the mean return equation apart from the case of Mexico, which was highly significant. The conditional variance equation, however, offers a very different narrative. The partisan variable was significant at varying levels of significance i.e., ranging between the 10% and 1% levels, for all pairs. However, the sign of the coefficient varied between nations.

For Canada, Japan, and Mexico, the partisan coefficient was negative, suggesting that periods of a Democrat presidency generally led to lower periods of volatility. This contrasts with the conclusions of Lobo and Tufte (1998) and Blomberg and Hess (1997). However, for the U.K. and the eurozone, the positive sign suggests higher periods of volatility during a Democrat presidency, which is consistent with what was expected. Even though we do not find a consistent impact of a Democrat presidency on volatility, we can argue that there is evidence that partisan factors do have some impact on the conditional variance, thus supporting the idea that the partisan cycle is a key variable.

The divided government variable as per Lohmann and O'Halloran (1994) suggests that we will see higher volatility for exchange rates during periods of divided government. We should note that the impact of this variable has not been investigated in the extant literature. It is expected that during a divided government, we see higher levels of isolationism and hence lower levels of trade openness from tighter trade policies. Hau (2002) and Mpofu (2016) have demonstrated an inverse relationship between trade openness and currency volatility, so we expect higher volatility during divided government, which is consistent with our results from our trade openness control variable. As expected, apart from in the case of the U.K. and the eurozone, the divided government variable was insignificant in the mean return equation. Turning to the conditional variance equation, we find some mixed, yet important, results. For the CAD and PESO, the findings were positive and consistently significant across most of the models. On a stand-alone basis, these results loosely support our hypothesis that we see higher levels of volatility during a divided government.

Interestingly, Japan tells a similar story with the divided government variable being positive and significant at the 1% level across all four models. The size of the coefficients suggests that the JPY experiences the highest volatility during a divided government. However, we find consistency in the coefficients for models 2, 3, and 4, which still confirms this conclusion compared to Canada.

In the U.K. and eurozone estimations, divided government was found to be insignificant across all the models, but it did display a positive coefficient, which is in line with our expectations. Generally, we can say that our theoretical hypothesis of higher volatility is a valid conclusion to draw from the results, as 3 out of 5 countries confirmed this.

Although our study mainly focuses on political variables, important conclusions can be drawn from the control variables in the conditional variance equations. Across most models, we found that inflation and crude oil were generally significant determinants of volatility. However, unexpectedly, we provide evidence of lower periods of volatility during higher volatility in the crude oil market, which is a surprising finding.

Generally, our findings were consistent with those of Hau (2002) and later literature, such as Stancik (2007) and Mpofu (2016). We found that apart from in the case of Mexico, the U.S. trade openness variable was significant with a negative coefficient. The latter confirms our assumption that higher trade openness potentially leads to lower volatility in the exchange rate. This is almost certain for the GBP and EURO, which displayed significance at the 1% level across most of the estimated models, possibly reflecting the strong economic relationship between these countries.

Japan displays limited significance for U.S. trade openness as only in model 1 did we find trade openness to be significant at the 5% level. Arguably, the later models found more consistency in the coefficient, indicating a better fit, thus making it generally challenging to conclude about the JPY. Furthermore, given that U.S. trade openness is relevant to the theoretical underpinning for the divided government variable, the lack of significance questions our understanding of the results presented by the divided government variable for Japan.

There are definitive signs of ARCH and GARCH effects in all five currencies and models. We see that across all currencies and varying models, both ARCH and GARCH terms are significant and positive at the 1% level. This is expected and is consistent with the earlier literature using the GARCH (1,1) model (Liu & Pauwels, 2012). The data from the ARCH term across the four models suggests that Canada experiences the highest level of volatility in the short run, with Japan experiencing the lowest levels of volatility. In the GARCH term, Japan experiences the greatest persistence in volatility based on a single time lag. The finding that the JPY experiences higher persistence than the CAD, GBP, PESO, and EURO are consistent with the findings of Lobo and Tufte (1998). However, as observed in the literature, there are often asymmetric effects whereby adverse news has a greater effect on volatility than good news[9].

In sum, as far as the electoral variables are concerned, the 6-month and 3-month variables were found to be insignificant. However, for the five currencies we did find the one-month electoral variable to be significant (albeit at 10%) and positive, suggesting an increase in volatility in the month before an election as uncertainty grows and the electionbecomes more contentious. Furthermore, the most important results came from the partisan and divided government variables. Generally, we see that there are partisan effectson exchange rates, albeit with variability in the sign of these coefficients, thus suggesting a varying impact on different currencies. Additionally, we find that divided government was a factor in increasing exchange rate volatility in three out of the five currencies, which helps to confirm our hypothesis as based on Lohmann and O'Halloran's divided government theory.

4.1 Discussion

In placing the obtained results into context with our initial theoretical stance, several points can be made. Although exchange rate volatility around election

9 It seems that the GARCH (1,1) model does not differentiate between bad and good news volatility, which is arguably a weakness.

cycles has been extensively investigated, the results have been inconsistent. However, there is a general consensus concerning an increase in volatility in the period before elections. Generally, our results are consistent with these prior findings, even though our election variable presented the weakest results of all the political variables. A solid conclusion we can draw is that exchange rate volatility increased one month prior to an election, and this was only significant at the 10% level across all models. This loosely supports the PBC hypothesis, whereby a potential rationale for the increase in volatility prior to election periods results from politicians attempting to target popular economic policies that will increase their chances of election, as distinct from their policies in non-election years.

Arguably, there is evidence that our results support a more straightforward explanation as presented by Garfinkelet al. (1999), who argue that volatility is related to uncertainty over the change in policy outcomes following an election, similar to the argument imparted by Freeman et al. (2000). This is more reminiscent of the partisan theory (PT), which states that in a close election, there is higher uncertainty regarding policy due to parties having different ideological positions, which is reflected in their policies. This explanation is not consistent with the assumption laid out in the PBC where an electorate is a homogenous group. In theory, economic policy should not vary because of ideological factors but due to an election cycle (Nordhaus, 1975). Therefore, PT is a more plausible explanation.

Furthermore, the insignificance of our three-month and six-month variables is not in line with broader evidence offered by Lobo and Tufte (1998), Siokis and Kapopoulos (2003), and Leblang and Bernhard (2006), whose variables were designed for an electoral year or 16 weeks before an election, hence suggesting that we will see higher volatility during these periods. An argument can be made that the insignificance of these variables can be explained by previous studies suggesting that not every election has the same significant implications.

For instance, Leblang and Bernhard (2006) found only 11 out of 23 elections in their sample to be statistically significant. This reflects the efficient market hypothesis theory, according to which only uncertainty will cause changes in the market environment, otherwise it is priced in. Arguably, the decision to take the electoral variable within a single aggregated study diluted our understanding of characteristics surrounding individual elections. Some elections were potentially more volatile than others and our results were too general. One of the problems of looking at partisan variables and electoral variables is that one is best suited to a disaggregated event study but the other needs time series analysis. This study was conducted using an aggregated time-series model,which may be why we obtained more robust results for the partisan variable and fragmented results for the electoral variables.

Our partisan variable turned out to be significant in the conditional variance. However, the results suggesting how the partisan cycle affects volatility are somewhat unclear, but based on three out of the five currencies exhibiting lower volatility, there is evidence for a loose conclusion surrounding Democrat presidencies experiencing lower volatility. In the foundational work by Lobo and Tufte (1998), they concluded that partisanship was a significant variable leading to higher volatility during a Democrat presidency, which was also consistent with Blomberg and Hess (1996).

There are several ways one can explain the conflicting evidence. Firstly, Lobo and Tufte's sample contained just one Democrat presidency, that of Jimmy Carter, which only lasted a single term, so merely four out of twenty years were Democrat-oriented, with their sample being overwhelmingly dominated by Republican presidencies. In contrast, our sample saw parts of three Democrat presidencies, which accounted for nine out of twenty years of our sample. Our larger sample might thus help to provide a better estimation of partisan effects. Additionally, the Carter presidency was fraughtwith economic shocks, such as stagflation and an oil crisis, which meant policy focus largelymoved away from full employment to anti-inflationary measures (Ponder, 2003). These policies are contrary to traditional Democrat policies, which we saw more of during the Clinton and Obama presidencies in the form of large spending packages and expansion of the welfare state. Moreover, for the most part, Obama's presidency covered a period of relative recovery after the global financial crisis.

Generally, the distinction between the levels of volatility between a Republican and Democrat presidency largely confirms the partisan theory. An argument can be made that this volatility reflects variations in policies along partisan and ideological lines resulting in the targeting of different optimal policies in accordance with the Philips curve trade-off, although, given the Democrats' traditional preference towards social welfare policy, we would expect an increase in volatility as traditionally pro-inflationary policies are implemented. However, there is an argument that partisan distinctions between traditional inflationary vs employment have become convoluted in the post-global financial crisis world. Pro-employment policies were the core of Trump's government, which pressured the Fed into cutting ratesto record lows to counteract the effects of protectionist policies and stimulate the economy (Smialek, 2019). Arguably, evidence from our trade openness and divided government variables support the conclusion that protectionist policies are associated with currency volatility. The fact that the results of our study are different to those of prior studies might suggest that partisan compositions and policy preferences have changed since the original studies of the 1990s.

The results of the divided government variable largely supported our hypothesis laid out by divided government theory (Lohmann & O'Halloran, 1994), suggesting that divided government can lead to higher volatility. The explanation put forward by this theory argues that this is dueto more isolationist policies being adopted as the executive is limited in its function in supporting foreign trade. Generally, the finding regarding higher trade openness lowering exchange rate volatility supports this notion. Although this theory is only a potential explanation, we find evidence of a connection between the two, and our findings generally add credence to Lohmann and O'Halloran (1994).

5. CONCLUDING REMARKS

In this paper, we sought to link U.S. domestic politics with exchange rate behaviour in a modern setting. We also sought to connect the dots between the literature on the international political economy with the finance literature. This is achieved by employing the GARCH methodology and utilising past literature and theoretical underpinnings to establish potential variables that can be used in further explaining exchange rate behaviour.

This study's unique direction is reflected by adding a previously uninvestigated variable (divided government) in the context of the international political economy literature. The political landscape has undoubtedly changed since the foundational literature of Lobo and Tufte (1998),Siokis and Kapopoulos (2003), and Leblang and Bernhard (2006).

Our investigation found mixed but interesting results. For one thing, the mean equation didnot yield significant results, which was expected and consistent with prior literature. Going deeper, the conditional variance equation, which focused on volatility, provided evidence that domestic U.S. politics has a significant effect on exchange rate volatility.

The most robust results we found were those for the partisan variable. Specifically, we found that partisanshipwas a significant variable for exchange rate return volatility across all currency pairs at a reasonable and consistent significance level. Furthermore, we found that Democrat presidencies present lower volatility than Republican ones, although this was not consistent across all the currency pairs; the evidence loosely suggests this to be the case. Such evidence broadly supports the notion put forward that ideological preferences in U.S. politics affectexchange rate volatility.

We also provide evidence to support the divided government theory. We find that there is a significant increase in volatility in relation to the CAD, PESO, and JPY during periods of divided government.We interpret this as evidence for the notion that periods of divided government see higher levels of isolationist policies by the U.S. as Congress limits the foreign policy role of the executive. Therefore, trade policy becomes more uncertain, and the U.S. experiences reduced trade openness. The evidence from the analysis of trade openness provides some support for this, i.e., the divided government theory of Lohmann and O'Halloran (1996). The latter was a previously unexplored variable in the study of foreign exchange markets. The relatively concrete results contribute to our understanding of politics and exchange rate behaviour.

We found limited evidence for the electoral variables consistent with the partisan business cycle theory. The most concrete evidence we report suggests a significant increase in volatility one month before an election across the five currencies. We argue that our results are more supportive of the partisan theory. Although this support is relatively weak, it is still valuable. On balance, we find that U.S. politics, through its various factors, influences global exchange rate behaviour, which might be a consequence of the hegemonic position of the U.S.

REFERENCES

Alesina, A. (1988). Macroeconomics and Politics. *NBER Macroeconomics Annual*, *3* (1), 13–62.

Alesina, A., & Roubini, N. (1992). Political Cycles in OECD Economies. *The Review of Economic Studies*, *59*(4), 663–688, https://doi.org/10.2307/2297992

Alesina, A., & Sachs, J. (1988). Political Parties and the Business Cycle in the United States. 1948-1984, *Journal of Money, Credit and Banking*, *20*(1), 63–82, https://doi.org/10.2307/1992667

Bachman, D. (1992). The effect of political risk on the forward exchange bias: the case of elections. *Journal of International Money and Finance*, *11*(2), 208–219, https://doi.org/10.1016/0261-5606(92)90042-V

Hofmann, B., & Park, T. (2020). The broad dollar exchange rate as an EME risk factor. Available at SSRN 4051584

Bekaert, G., Harvey, C.R., Lundblad, C.T., & Siegel, S. (2014). Political risk spreads. *Journal of International Business Studies*, *45*(1), 471–493, https://doi.org/10.1057/jibs.2014.4

Blomberg, S.B., & Hess, G.D. (1997). Politics and exchange rate forecasts. *Journal of International Economics*, *43*(1–2), 189–205, https://doi.org/10.1016/S0022-1996(96)01466-3

Bouraoui, T., & Hammami, H. (2017). Does political instability affect exchange rates in Arab Spring countries?. Applied Economics, 49(55), 5627-5637

Brogaard, J., Dai, L., Ngo, P.T., & Zhang, B. (2020). Global Political Uncertainty and Asset Prices. *Review of Financial Studies*, *33*(4), 1737–1780, https://dx.doi.org/10.2139/ssrn.2488820

dos Santos, M.B.C., Klotzle, M.C., & Pinto, A.C.F. (2021). The impact of political risk on the currencies of emerging markets. *Research in International Business and Finance*, *56*, 101375. https://doi.org/10.1016/j.ribaf.2020.101375

Filippou, I., Gozluklu, A.E., & Taylor, M. (2018). Global Political Risk and Currency Momentum. *Journal of Financial and Quantitative Analysis, 53*(5), 2227–2259. https://doi.org/10.1017/S0022109018000686

Freeman, J.R., Hays, J. C., & Stix, H. (2000). Democracy and Markets: The Case of Exchange Rates. *American Journal of Political Science*, *44*(3), 449–468. https://doi.org/10.2307/2669258

Garfinkel, M.R., Glazer, A., & Lee, J. (1999). Election Surprises and Exchange Rate Uncertainty. *Economics and Politics*, *11*(3), 255–274, https://doi.org/10.1111/1468-0343.00061

Hau, H. (2002). Real Exchange Rate Volatility and Economic Openness: Theory and Evidence. *Journal of Money, Credit and Banking*, *34*(3), 611–630. http://www.jstor.org/stable/3270734

Hibbs Jr, D.A. (1994). The Partisan Model of Macroeconomic Cycles: More Theory and Evidence for The United States. *Economics and Politics*, 6(1), 1–23, https://doi.org/10.1111/j.1468-0343.1994.tb00081.x

Junttila, J., & Korhonen, M. (2012). The role of inflation regime in the exchange rate pass-through to import prices. *International Review of Economics and Finance*, *24*(1), 88–96, https://doi.org/10.1016/j.iref.2012.01.005

Kutan, A.M., & Zhou, S. (1995). Sociopolitical instability, volatility, and the bid-ask spread: Evidence from the free market for dollars in Poland. *Open Economics Review*, *6*(6), 225–236, https://doi.org/10.1007/BF01000082

Leblang, D., & Bernhard, W. (2006). Parliamentary Politics and Foreign Exchange Markets: The World According to GARCH. *International Studies Quarterly*, *50*(1), 69-92 https://www.jstor.org/stable/3693552

Liu, L. G., & Pauwels, L.L. (2012). Do External Political Pressures Affect the Renminbi Exchange Rate? *Journal of International Money and Finance*, *31*(6), 1800–1818, https://doi.org/10.1016/j.jimonfin.2012.04.001

Lobo, B.J., & Tufte, D. (1998). Exchange rate volatility: Does politics matter? *Journal of Macroeconomics*, *20*(2), 361–365, https://doi.org/10.2139/ssrn.4197322

Lohmann, S., & O'Halloran, S. (1994). Divided government and U.S. trade policy: Theory and evidence. *International Organization*, *48(*4), 595–632, https://www.jstor.org/stable/2706897

Melvin, M., & Tan, K.H. (1996). Foreign Exchange Market Bid-Ask Spreads and the Market Price of Social Unrest. *Oxford Economic Papers*, *48*(2), 329–341.

Menkhoff, L., Sarno, L., Schmeling, M., & Schrimpf, A. (2017). Currency Value. *The Review of Financial Studies*, *30*(2), 416–441. http://www.jstor.org/stable/26142245

Menkhoff, L., Sarno, L., Schmeling, M., & Schrimpf, A. (2012). Carry Trades and Global Foreign Exchange Volatility. *The Journal of Finance*, *67*(2), 681–718. http://www.jstor.org/stable/41419708

Mpofu,T.R. (2016). *The Determinants of Exchange Rate Volatility in South Africa*. (Economic Research Southern Africa Working Paper 604).

Neeley, C.J (2005). An Analysis of Recent Studies of the Effect of Foreign Exchange Intervention. *Federal Reserve Bank of St. Louis Review*, *87*(6), 685–717.

Nordhaus,W.D. (1975). The Political Business Cycle. *The Review of Economic Studies*, *42*(2), 169–190, https://doi.org/10.2307/2296528

Ohmae, K.I. (1995). *The End of the Nation State: The Rise of Regional Economies*. Free press.

Ponder, D.E. (2003). [Review: Jimmy Carter's Economy: Policy in an Age of Limits by W. C. Biven]. *Presidential Studies Quarterly*, *33*(2), 459–461.

Proti, N.P. (2013). Exchange rate fluctuations-shock in Tanzania An empirical analysis. *Scholarly Journal of Business Administration*, *3*(1), 12–19.

Saeed, A., Awan, R., Sial, M., & Sher, F. (2012) An Econometric Analysis of the Determinants of Exchange Rate in Pakistan. *International Journal of Business and Social Science, 3*(6), 184–196.

Siokis, F., & Kapopoulos, P. (2003). Electoral management, political risk and exchange rate dynamics: the Greek experience. *Applied Financial Economics, 13*(4), 279–285, https://doi.org/10.1080/13504850210128839

Smialek, J. (2019). Federal Reserve Cuts Interest Rates for Third Time in 2019. https://www.nytimes.com/2019/10/30/business/economy/federal-reserve-interest-rates.html

Stancik, J. (2007). Determinants of Exchange-Rate Volatility: The Case of the New E.U. Members. *Czech Journal of Economics and Finance*, *57*(9), 414–432.

Vogiazas, S., Alexiou, C., & Ogan, O. (2019) Drivers of real effective exchange rates in high and upper-middle income countries. *Australian Economic Papers*, *58*(1). 41–53, https://doi.org/10.1111/1467-8454.12139

Vortelinos, D.I., & Saha, S. (2016). The impact of political risk on return, volatility and discontinuity: Evidence from the international stock and foreign exchange markets. *Finance Research Letters*, *17*(1), 222–226, https://doi.org/10.1016/j.frl.2016.03.017

WITS (2021a). Canada trade summary. World Integrated Trade Solution. https://wits.worldbank.org/CountryProfile/en/Country/CAN/Year/LTST/Summary. Last accessed: 9th October 2021

WITS. (2021b). United Kingdom trade summary. World Integrated Trade Solution. https://wits.worldbank.org/CountryProfile/en/Country/GBR/Year/2019/Summary.

WITS. (2021c). Japan trade summary. World Integrated Trade Solution. https://wits.worldbank.org/CountryProfile/en/Country/JPN/Year/2019/Summary.

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APPENDIX

Table A. Definition of variables and sources

Political Variables		
U.S. election 1 month	An indicator variable equal to one in a month that is one month prior to a U.S. Presidential election	Federal Election Commission
U.S. election 3 month	An indicator variable equal to one in a month that is three months prior to a U.S. Presidential election	Federal Election Commission
U.S. election 6 month	An indicator variable equal to one in a month that is six months prior to a U.S. Presidential election	Federal Election Commission
Divided Government	A dummy variable which assumes the value of 1 during periods of divided government and 0 otherwise	Brooking
Partisan	A dummy variable assuming the value of 1 during a Democrat presidency and 0 during a Republican presidency	Authors
Macroeconomic Controls		
Recession	OECD recession indicator	OECD
Trade Openness	Trade in goods and services: Import + Exports (% GDP)	OECD
Inflation	Consumer Price Index	OECD
Crude Oil	Natural logarithm of WTI Crude Oil price	FRED
Money Supply (M2)	Growth rate of the Money Supply (M2)	OECD

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FINANCIAL INCLUSION, MACROPRUDENTIAL POLICY, AND CRISIS

ABSTRACT: *The study examines the role played by financial inclusion (FI) and macroprudential policy (MPP) to prevent financial crisis or reduce the severity of a financial crisis going forward using a panel of 138 countries covering the years 2004–2017. To attain these objectives through robust experimentation and support policy formulation, we employ aggregated measures of FI and MPP and use advanced econometric models. Our findings show that, although FI initially decreases the likelihood of a crisis, the probability of a crisis increases after a certain level of inclusion is reached. In contrast, countries with MPP are less likely to have a crisis than countries without MPP. Furthermore, the FI-MPP interaction complements itself and plays a vital role in reducing the likelihood of a crisis. Our results are robust and could be useful for policymakers to formulate policies in order to prevent a crisis or reduce its severity going forward.*

KEY WORDS: *Crisis, Financial inclusion index, Macroprudential policy.*

JEL CLASSIFICATION: G18, G21, G28

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1. INTRODUCTION

Broadening inclusive financial services to the poor and underprivileged segments of society is considered one of the most important catalysts to reduce extreme poverty, enhance mutual prosperity, foster integrity, and attain several Sustainable Development Goals. Policymakers and international organisations, such as the International Monetary Fund, the Group of Twenty (G20), the Alliance for Financial Inclusion (AFI), and the Consultative Group to Assist the Poor (CGAP), have implemented many programmes in both developing and developed countries to improve access to formal finance. In spite of numerous policy initiatives, a recent survey by the World Bank shows that around 2.5 billion adults, more than half the adults around the globe, are excluded from formal channels of access to finance. Moreover, there is wide disparity of FI among countries, with over seventy-two per cent of adults being excluded from conventional financial services in developing countries, while only nineteen percent are excluded in developed economies.

FI narrows the income gap and reduces poverty (Bruhn & Love, 2014), promotes retail deposits (Allen et al., 2016), favours schooling and learning (Flug et al., 1998), creates jobs (Prasad, 2010), enhances happiness and psychological health (Angelucci et al., 2013), boosts the establishment of new business organisations (Klapper et al., 2006; Banerjee et al., 2013), assists efficient decision making (Mani et al., 2013), encourages entrepreneurs to take risk and invest more (Cumming et al., 2014), and decreases the asymmetry of information between borrowers and lenders (Petersen & Rajan, 1995). Several leading studies have already established a robust association between access to formal financial services and both economic growth as well as development (Demirguc-Kunt & Klapper, 2012; Honohan, 2004; Sahay et al., 2015a, 2015b). However, greater financial access may also induce the possibility of crisis (Torre et al., 2012; Sahay et al., 2015a). The 2008–2009 global financial crisis is an example which shows that ambitious initiatives inspiring larger access to finance may result in unintended consequences (McLean & Nocera, 2010). Although FI may lead to broader access and a more diversified base of deposits, lower information asymmetry, greater diversification opportunity, economies of scale, and competitive advantage, hazards might arise from the resultant hasty credit expansion of new institutions and instruments for promoting FI.

Furthermore, MPP, a buzz phrasenowadays, has regained its importance among regulators and economists since the 2008–2009 global financial crisis. The primary goal of these rules is to reduce systemic risk by limiting credit expansion, guarantee financial sector stability, and reduce the chance of a crisis. Many studies, such as Cerutti et al. (2017), Fendoğlu (2017), and Claessens et al. (2015)have been undertaken to explore the effectiveness of MPP in reducing credit growth, which is an intermediate goal. The ultimate objective of devising and implementing MPP tools is to avoid or overcome the risk of financial crises, yet there is a dearth of studies that analyse the efficacy of MPP with respect to this ultimate goal. Moreover, to the best of our knowledge, no study effort has yet been undertaken to evaluate the role of MPP in the context of extending FI to avert the occurrence of crises. With this in mind, using an unbalanced panel of 138 nations from 2004 to 2017, we analyse the joint impact of FI and MPP in reducing the chance of a crisis. We use both logit and probit models in our estimations to achieve a comprehensive understanding and robustness, and to facilitate policy implications.

The results reveal that, although FI initially decreases the likelihood of a crisis, the probability of a crisis increases after achieving a certain level of inclusion. Countries with MPP are less likely to have crises than countries with no MPP. Moreover, the FI-MPP interaction complementarily reduces the propensity of a crisis.

The contribution of the paper is manifold. First, it extends the existing literature by examining the influence of MPP in preventing financial crises, which is the ultimate goal. Second, it also explores the role of FI in averting financial catastrophes. Lastly, it analyses the combined impact of MPP and FI in preventing or reducing the propensity of crises.

The rest of this paper is organised as follows: Section 2 describes the data and methods, Section 3 displays the results and discusses them, and Section 4 concludes.

2. DATA AND METHODS

2.1. Data

To examine the combined impact of FI and MPP on the likelihood of a crisis, we utilise unbalanced panel data from 138 countries over the years 2004-2017 collected from various sources. We chose this timeframe because FI data is only available for this period. Data for the FI index are derived from the International Monetary Fund's FAS database (IMF, 2018); MPP data are sourced from Cerutti et al. (2017). WGI, WDI, and GFDD databases are used to gather crisis data as well as all control factors (World Bank, 2019a; World Bank, 2019b; World Bank, 2019c).

2.2. Financial inclusion

FI is a basic concept that may be described as the provision of formal financial services to everyone, especially low-income and disadvantaged groups, on reasonable terms and conditions.

Khan (2011) defines financial inclusion as "… *the process of ensuring access to financial services and timely and adequate credit where needed by vulnerable groups such as weaker sections and low income groups at an affordable cost. It primarily represents access to a bank account backed by deposit insurance, access to affordable credit and the payments system.*"

To quantify FI, we created the financial outreach index (FOI) using principal component analysis (PCA) to examine its overall impact. Because multiple FI indicators are strongly associated, substitutable, or complimentary in nature, including these indicators in a single model at the same time may give incorrect findings. To address the issues of over-parameterisation and multicollinearity, we employ four access indicators: the number of bank branches per 100,000 people, the number of ATMs per 100,000 people, the number of bank branches per 1000 square kilometres, and the number of ATMs per 1000 square kilometres to create the FOI using the PCA. The intuition for choosing these specific FI indicators is that they capture the distribution of financial services across the population as well as geographic areas.

Before conducting the PCA, all indicators are adjusted using the min-max normalisation procedure (1).

$$nmx = \frac{X_i - X_{min}}{X_{max} - X_{min}} \quad (1)$$

where X_{min} = minimum data point and X_{max} = maximum data point.

The eigenvalues of the four FOI components are 2.06, 1.35, 0.44, and 0.14, indicating that the first component explains 52% of the variation in the four indicators. We use the first component to build the index using the following equation (2).

$$\text{Financial Outreach Index} = \sum_{i=1}^{n} \gamma_{ij} x_i \quad (2)$$

γ_{ij} are the loadings of components or weights derived from PCA, and x_i are the original variables. A higher value of this index means greater financial access in an economy.

The Kaiser-Meyer-Olkin (KMO) measures of sampling adequacy are 0.61, and the p values for the Bartlett's test of sphericity are less than the 0.01 significance level, indicating that the variables used in the PCA are appropriate.

2.3. Macroprudential regulations

Though the term macroprudential was first used by the Euro Currency Standing Committee in their research on multinational disbursement of bank loans at the Bank for International Settlements (BIS) Committee during the late 1970s (Clement, 2010), it reemerged as a policy concern after the 2008–2009 global financial crisis. This recent crisis has led to criticism of the established doctrine, long-held convictions, and hypotheses on financial systems' risk regulation and highlighted the importance of a macroprudential perspective of regulation and surveillance for ensuring financial stability.

There are several sources for measuring MPP, such as Lim et al. (2011), Shim et al. (2013), Chantapacdepong and Shim (2014), Boh et al. (2017), and others. The majority of these macroprudential datasets span a relatively short period and cross sections, for example, Lim et al. (2011) include only 38 nations. We chose the 2018 version of the Cerutti et al. (2017) macroprudential policy dataset because of its comprehensive coverage. This dataset covers the use of twelve macroprudential measures in 160 countries from 2000 to 2017.

Table 1 Definition of Macroprudential Policy Instruments

Instrument	Short form	Definition
Survey Instruments		
Loan-to-Value Ratio	LTV	"Constrains highly levered mortgage down payments by enforcing or encouraging a limit or by determining regulatory risk weights."
Debt-to-Income Ratio	DTI	"Constrains household indebtedness by enforcing or encouraging a limit."
Time-Varying/Dynamic Loan-Loss Provisioning	DP	"Requires banks to hold more loan-loss provisions during upturns."
General Countercyclical Capital Buffer/Requirements	CTC	"Requires banks to hold more capital during upturns."
Leverage Ratio	LEV	"Limits banks from exceeding a fixed minimum leverage ratio."
Capital Surcharges on SIFIs	SIFI	"Requires Systemically Important Financial Institutions to hold a higher capital level than other financial institutions"
Limits on Interbank Exposures	INTER	"Limits the fraction of liabilities held by the banking sector or by individual banks."
Concentration Limits	CONC	"Limits the fraction of assets held by a limited number of borrowers"
Limits on Foreign Currency Loans	FC	"Reduces vulnerability to foreign-currency risks."
Reserve Requirement Ratios	RR	"Limits credit growth; can also be targeted to limit foreign-currency credit growth"
Limits on Domestic Currency Loan	CG	"Limits credit growth directly."
Levy/Tax on Financial Institutions	TAX	"Taxes revenues of financial institutions."
Derived Instruments		
Loan-to-Value Ratio Caps	LTV_CAP	"Restricts to LTV used as a strictly enforced cap on new loans, as opposed to a supervisory guideline or merely a determinant of risk weights."
FX and/or Countercyclical Reserve Requirements	RR_REV	"Restricts to RR which i) imposes a wedge of on foreign currency deposits (as determined by the answer to question 9.1.4.2 "Please specify the level of reserve requirements applied to specific bases identified in the question above on the last day of the year preceding the submission of this survey"), or ii) is adjusted countercyclically (as determined by the answer to the question 9.1.8 "Please specify whether this tool is intended to be adjusted"

Source: *Cerutti et al. (2017)*

The instruments are General Countercyclical Capital Buffer/Requirement (CTC); Leverage Ratio for banks (LEV); Time-Varying/Dynamic Loan-Loss Provisioning (DP); Loan-to-Value Ratio (LTV); Debt-to-Income Ratio (DTI); Limits on Domestic Currency Loans (CG); Limits on Foreign Currency Loans (FC); Reserve Requirement Ratios (RR); Levy/Tax on Financial Institutions (TAX); Capital Surcharges on SIFIs (SIFI); Limits on Interbank Exposures (INTER); and Concentration Limits (CONC). The dataset is developed using simple binary measurements of whether or not the instruments were found in a certain country-year. Following compilation, a composite macroprudential index (MPI) is created by adding the scores of all 12 policies. Table 1 shows the instrument definitions used by Cerutti et al. (2017).

2.4. Banking crisis

To represent the banking crisis, we utilise a dummy variable. A banking crisis, according to the World Bank, begins with significant signs of financial distress and policy intervention in response to significant losses in the banking system and ends before the year in which both real GDP growth and real credit growth are positive for at least two consecutive years.

2.5. Control variables

Choosing a control variable is critical for any empirical research, especially for a cross-country analysis because there may be a lot of variability and unobserved variables. Furthermore, failing to include a confounder and to exclude a collider variable as a control variable can lead to major errors in inferences (Rohrer, 2018).

As a result, we apply multiple control variables based on the current literature to control the country-specific heterogeneity of our sample. Following Sahay et al. (2015a, b), trade globalisation (TG), the financial development index (FID), and bank concentration (CON) are used to control the banking industry's trade openness, financial development, and market structure. Financial freedom (FF) is a term used to describe how different nations' banks are efficient and independent.

We employ the growth rate of GDP per capita (gGDPC) and total GDP to account for economic growth and development, following Bermpei et al. (2018). The efficacy of MPP is determined by their interaction with monetary policy (Bruno et al., 2017; Akinci & Olmstead-Rumsey, 2015; Claessens, 2015), and monetary policy influences financial stability (Angeloni et al. 2015). Thus, in line with Dutta and Saha (2021), the consumer price index (CPI) is employed to account for the variance in monetary policy among our sample's nations. Furthermore, total population (POP) is used to regulate a country's market size, which is similar to Saha and Dutta (2022).

2.6. Model

To investigate whether FI and MPP prevent a crisis or reduce the severity of a crisis going forward, we use logistic and probit models. Regardless of the fact that these methods are both symmetric binary choice models, the dissimilarity between these two methods lies in this postulation regarding the distribution of the error-term. The logistic model presumes a standard logistic distribution of error-terms, whereas the probit model presumes a normal distribution of error-terms. For this analysis, our dependent variable is crisis, which is a dummy variable representing the occurrence of a banking crisis expressed by value 0 (no crisis) and 1 (crisis). We will use the following equations to quantify the relationship between the explanatory variables and the likelihood of a crisis. Equations (3) and (4) represent the logit and probit models, respectively.

$$ln\left(\frac{p_i}{(1-p_i)}\right) = \sum_{k=0}^{k=n} \beta_k x_{ik} \tag{3}$$

where P_i is the probability of a crisis , P_0 = 1- P_i is the probability of no banking crisis with the probability ratio P_i / P_0, and ln represents the logic transformation. β are values of coefficients estimated from the data set by maximising the log-likelihood function. x_{ik} is a set of {k} explanatory variables used to predict the probability of a banking crisis.

The probit regression is a specialised regression model of binomial response variables and is also used to analyse the relationship between binary dependent and explanatory variables.

$$\Phi^{-1}(p_i) = \sum_{k=0}^{k=n} \beta_k x_{ik} \tag{4}$$

3. RESULT AND DISCUSSION

3.1. Summary statistics

Table 2 displays the descriptive statistics for the data used in our investigation. To decrease the impact of outliers, all variables except unitary and dummy variables are winsorised at the 1st and 99th percentile levels. The FOI standard deviations show significant cross-country variance. The MPI's minimal value of 0 to its highest value of 10 also shows a wide range of countries' adoption of macroprudential legislation. Because our analysis covers nations with various income categories and growth regimes, the GDP and gGDPC exhibit a greater range. All other characteristics vary among nations as well. Crisis is a dummy variable with a value of 1 representing a banking crisis and a value of 0 indicating the absence of a crisis.

Table 2 Descriptive statistics

Variable	Obs	Mean	Std.Dev.	Min	Max
FOI	1916	0.106	0.13	0	1
MPI	2184	2.11	1.75	0	10
GOV	2164	0.493	0.228	0	1
CON	1938	66.852	19.011	28.713	100
FF	2025	52.533	18.673	0	90
TG	2156	56.695	18.905	8.804	99.551
GDP (in billions)	2181	301.85	779.19	0.17	4900
gGDPC	2171	2.57	3.78	-12.209	14.746
CPI	2172	103.769	27.53	50.637	243.97
FID	2100	0.285	0.269	0.002	1
POP (in millions)	2184	40.32	130.1195	0.057	1100
CRISIS	2184	0.051	0.22	0	1

3.2. Effectiveness of financial inclusion and macroprudential policy to prevent crisis

To investigate whether FI and MPP are effective at preventing a crisis, we estimate the equations using logit and probit regressions. By estimating both models, we aim to study if there are any relevant differences between them in the obtained results. The results reported in Table 3 show the findings are almost identical.

The coefficient of lag FOI is negatively significant, whereas the quadratic lag term of FOI is positively significant, suggesting that FI initially decreases the likelihood of a crisis; however, after achieving a certain level of inclusion, the probability of a crisis increases. The following equations (5) and (6) are written from the results of the probit and logit average marginal effect analyses.

$$\mathrm{Crisis} = -12.932\mathrm{FOI}_{t-1} + 25.205\mathrm{FOI}_{t-1}^{2} \quad (5)$$

$$\mathrm{Crisis} = -24.919\mathrm{FOI}_{t-1} + 50.902\mathrm{FOI}_{t-1}^{2} \quad (6)$$

By constructing a scale of the FOI values starting from the minimum value (among countries and time) and increasing it by 0.1 to the maximum, we find 0.52 (0.49) is the level of threshold at which financial inclusion (the outreach dimension) increases the likelihood of a crisis in equation 5 (equation 6). These results should be interpreted with caution, as different dimensions of FI may have different probabilities in regard to crises. Countries with MPPs are 8.836 per cent (according to Model 5) less likely to have a crisis than countries with no MPPs. The coefficient of interaction term is negative and significant, implying that they complement each other to decrease the propensity of a crisis.

Therefore, the joint impact of FOI and MPI at different levels of FOI can be written as the following equation (7) and (8) as per probit and logit average marginal effect analyses.

$$\mathrm{Crisis} = -12.932\mathrm{FOI}_{t-1} + 25.205\mathrm{FOI}_{t-1}^{2} + (-8.836)\times\mathrm{MPI} + (-13.201)\times\mathrm{FOI} \quad (7)$$

$$\mathrm{Crisis} = -24.919\mathrm{FOI}_{t-1} + 50.902\mathrm{FOI}_{t-1}^{2} + (-17.048)\times\mathrm{MPI} + (-34.275)\times\mathrm{FOI} \quad (8)$$

By constructing a similar scale of the FOI and using different numbers of MPI, we find that at the 0.50 level of the FOI and 3 MPI, the likelihood of a crisis decreases by 33.27 per cent as per equation (7).

The classification accuracy of both probit and logit functions is very high and similar, 93.24% and 93.31%, respectively (Appendix Table A1).

Table 3 Baseline estimation: joint impact of financial inclusion and macroprudential policy

Variables	(1)	(2)	(3)	(4)	(5)	(6)
	Probit	Logit	Probit	Logit	Marginal effect (Probit)	Marginal effect (Logit)
FOI_{t-1}	-14.175***	-28.679***	-12.932***	-24.919***	-12.932***	-24.919***
	(2.920)	(5.814)	(3.051)	(6.033)	(3.051)	(6.033)
FOI^2_{t-1}	25.056***	52.433***	25.205***	50.912***	25.205***	50.912***
	(6.626)	(13.143)	(6.677)	(12.940)	(6.677)	(12.940)
MPI_{t-1}	1.485	2.804	0.249	-0.734	0.249	-0.734
	(1.563)	(3.009)	(1.766)	(3.512)	(1.766)	(3.512)
MPI^2_{t-1}	-8.407**	-15.992**	-8.836**	-17.048**	-8.836**	-17.048**
	(3.808)	(7.544)	(3.840)	(7.747)	(3.840)	(7.747)
$FOI_{t-1} \times MPI_{t-1}$			-13.201**	-34.275*	-13.201**	-34.275*
			(5.738)	(17.223)	(5.738)	(17.223)
CON_{t-1}	-0.522*	-1.002*	-0.559*	-1.074*	-0.559*	-1.074*
	(0.288)	(0.551)	(0.289)	(0.553)	(0.289)	(0.553)
FF_{t-1}	-0.429	-0.881	-0.545	-1.190	-0.545	-1.190
	(0.490)	(0.960)	(0.500)	(0.979)	(0.500)	(0.979)
TG_{t-1}	0.711*	1.246	0.872**	1.605**	0.872**	1.605**
	(0.400)	(0.768)	(0.419)	(0.798)	(0.419)	(0.798)
GDP_{t-1}	-0.827*	-1.754**	-0.822*	-1.723*	-0.822*	-1.723*
	(0.472)	(0.875)	(0.476)	(0.887)	(0.476)	(0.887)
$gGDPC_{t-1}$	-1.714***	-3.620***	-1.708***	-3.587***	-1.708***	-3.587***
	(0.433)	(0.824)	(0.434)	(0.828)	(0.434)	(0.828)
GOV_{t-1}	0.348	1.060	0.406	1.225	0.406	1.225
	(0.612)	(1.193)	(0.617)	(1.207)	(0.617)	(1.207)
CPI_{t-1}	-1.007	-2.331	-0.861	-1.852	-0.861	-1.852
	(0.746)	(1.748)	(0.748)	(1.744)	(0.748)	(1.744)
FID_{t-1}	0.359	0.762	0.380	0.821	0.380	0.821
	(0.402)	(0.777)	(0.403)	(0.776)	(0.403)	(0.776)
POP_{t-1}	1.403***	2.788***	1.446***	2.869***	1.446***	2.869***
	(0.508)	(0.970)	(0.515)	(0.988)	(0.515)	(0.988)
Constant	-1.677***	-3.059***	-1.570***	-2.732**	-1.570***	-2.732**
	(0.550)	(1.119)	(0.552)	(1.124)	(0.552)	(1.124)
Year dummy	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1,494	1,494	1,494	1,494	1,494	1,494

A robustness test with unwinsorised data is given in the Appendix. The results are also consistent with our main findings.

4. CONCLUSION

The ample literature on financial stability provides evidence of specific governmental attention to maintain stability, yet there is still a scarcity of research to investigate the elements influencing the likelihood of a financial crisis. There are few empirical studies that explore the influence of MPP on crises, and none of them take the role of FI into account while examining this nexus. Using an unbalanced panel of 138 nations from 2004 to 2017, we analysed how FI conditioned the efficacy of MPP on the risk of crises. We employed both logit and probit models to gain a comprehensive understanding, robustness, and to facilitate policy implications.

Our findings indicate that while FI increases the likelihood beyond a certain threshold, MPP is useful in avoiding the emergence of a crisis, and their interactions complementarily reduce the propensity of crisis. Specifically, a country that scores 0.50 on financial inclusion and has 3 MPI measures is 33.27 per cent less likely to be hit by a crisis.

The findings are important for developing suitable policies to lower the likelihood of a crisis and promote financial stability. Policymakers should encourage FI when it is below the threshold level and promote FI cautiously when it reaches the threshold level in order to prevent the occurrence of a crisis. Moreover, along with promoting FI, they should also devise and implement a macroprudential regulatory framework and rigorous supervision to minimise the probability of a crisis.

REFERENCES

Akinci, O., & Olmstead-Rumsey, J. (2015). *How effective are macroprudential policies? An Empirical investigation.* (International Finance Discussion Papers 1136). Board of Governors of the Federal Reserve System. http://dx.doi.org/10.17016/IFDP.2015.1136

Allen, F., Demirguc-Kunt, A., Klapper, L., & Peria, M.S.M. (2016). The foundations of financial inclusion: Understanding ownership and use of formal accounts. *Journal of Financial Intermediation, 27*, 1–30. https://doi.org/10.1016/j.jfi.2015.12.003

Angeloni, L., Faia, E., & Duca, M. L. (2015). Monetary policy and risk taking. *Journal of Economic Dynamics and Control, 52*, 285–307. https://doi.org/10.1016/j.jedc.2014.12.001

Angelucci, M., Karlan, D., & Zinman, J. (2013). *Win some lose some? Evidence from a randomized microcredit program placement experiment by Compartamos Banco* (NBER Working Paper No. 19119). The National Bureau of Economic Research. https://www.nber.org/papers/w19119

Banerjee, A., Chandrasekhar, A.G., Duflo, E., & Jackson, M.O.(2013) The diffusion of microfinance. *Science, 341*(6144), 1236498. https://doi.org/10.1126/science.1236498

Bermpei, T., Kalyvas, A., & Nguyen, T.C. (2018). Does institutional quality condition the effect of bank regulations and supervision on bank stability? Evidence from emerging and developing economies. *International Review of Financial Analysis, 59*, 255–275. https://doi.org/10.1016/j.irfa.2018.06.002

Boh, S., Borgioli, S., Coman, A.(B.), Chiriacescu, B., Koban, A., Kusmierczyk, P., Pirovano,M., Thomas Schepens, T., & Veiga, J. (2017). *European Macroprudential Database* (ECB Statistics Paper Series No. 32). European Central Bank. https://www.ecb.europa.eu/pub/pdf/scpsps/ecb.sps32~08c68f95ed.en.pdf

Bruhn, M., & Love, I. (2014). The real impact of improved access to finance: Evidence from Mexico. *The Journal of Finance, 69*(3), 1347–1376. https://doi.org/10.1111/jofi.12091

Bruno, V., Shim, I. and Shin, H. S. (2017). Comparative assessment of macroprudential policies. *Journal of Financial Stability, 28*, 183–202. https://doi.org/10.1016/j.jfs.2016.04.001

Cerutti, E., Claessens, S., & Laeven, L. (2017). The use and effectiveness of macroprudential policies: New evidence. *Journal of Financial Stability, 28*, 203–224. https://doi.org/10.1016/j.jfs.2015.10.004

Chantapacdepong, P., & Shim, I. (2014). *Correlations across Asia-Pacific bond markets and the impact of capital flow measures* (BIS Working Paper 472). Bank for International Settlements. https://www.bis.org/publ/work472.htm

Claessens, S. (2015). An Overview of Macroprudential Policy Tools. *Annual Review of Financial Economics, 7*, 397–422. https://doi.org/10.1146/annurev-financial-111914-041807

Claessens, S., Ghosh, S.R., & Mihet, R. (2014). *Macro-Prudential Policies to Mitigate Financial System Vulnerabilities.* (No. 14/155). International Money Fund. https://www.imf.org/external/pubs/ft/wp/2014/wp14155.pdf

Clement, P. (2010). *The term _macroprudential': Origins and evolution*. BIS Quarterly Review, March. Retrieved from https://www.bis.org/publ/qtrpdf/r_qt1003h.pdf

Cumming, D., Johan, S., & Zhang, M. (2014). The economic impact of entrepreneurship: Comparing international datasets. *Corporate Governance: An International Review, 22* (2), 162–178. https://doi.org/10.1111/corg.12058

De la Torre, A., Ize, A., & Schmukler, S. L. (2012). Financial Development in Latin America and the Caribbean: The Road Ahead. *World Bank Latin America and the Caribbean Studies*. http://documents.worldbank.org/curated/en/784141468053952645/pdf/657440REPLACEM0110USE0S

AME0BOX0INFO.pdfDemirguc-Kunt, A., & Klapper, L. (2012). Measuring Financial Inclusion: The Global Findex Database. *World bank policy research Working Paper*, (6025). https://doi.org/10.1596/1813-9450-6025

Dutta, K.D., & Saha, M. (2021). Nexus of governance, macroprudential policy and financial risk: cross-country evidence. *Economic Change and Restructuring* , 54, 1253–1298. https://doi.org/10.1007/s10644-020-09301-9

Fendoğlu, S. (2017). Credit cycles and capital flows: Effectiveness of the macroprudential policy framework in emerging market economies. *Journal of Banking & Finance, 79*, 110–128 https://doi.org/10.1016/j.jbankfin.2017.03.008

Flug, K., Spilimbergo, A., & Wachtenheim, E. (1998). Investment in education: do economic volatility and credit constraints matter? *Journal of Development Economics, 55(*2), 465–481. https://doi.org/10.1016/S0304-3878(98)00045-5

Honohan, P. (2004). Financial development, growth and poverty: How close are the links? In C.A.E. Goodhart (Eds.). *Financial development and economic growth*. London: Palgrave Macmillan. https://doi.org/10.1057/9780230374270_1

International Monetary Fund. (2018). Financial Access Survey (FAS) [Data file]. http://data.imf.org/?sk=E5DCAB7E-A5CA-4892-A6EA-598B5463A34C&sId=1496861011734

Khan, H. (2011) Financial Inclusion and Financial Stability: Are They Two Sides of the Same Coin? BANCON 2011, Indian Bankers Association and Indian Overseas Bank, Chennai. https://www.bis.org/review/r111229f.pdf

Klapper, L., Laeven, L., & Rajan, R. (2006). Entry regulation as a barrier to entrepreneurship. *Journal of Financial Economics, 82*(3), 591–629. https://doi.org/10.1016/j.jfineco.2005.09.006

Lim, C.H., Columba, F., Costa, A., Kongsamut, P., Otani, A., Saiyid, M., … & Wu, X. (2011). *Macroprudential policy: what instruments and how to use them? Lessons from country experiences*. (IMF Working Paper No. 11/238). https://www.imf.org/external/pubs/ft/wp/2011/wp11238.pdf

Mani, A., Mullainathan, S., Shafir, E., & Zhao, J. (2013). Poverty impedes cognitive function. *Science, 341*(6149), 976–980. https://doi.org/10.1126/science.1238041

Mclean, B., & Nocera, J. (2010). All the devils are here: the hidden history of the financial crisis. Portfolio

Petersen, M.A., & Rajan, R.G. (1995). The effect of credit market competition on lending relationships. *The Quarterly Journal of Economics, 110*(2), 407–443. https://doi.org/10.2307/2118445

Prasad, E.S. (2010). *Financial sector regulation and reforms in emerging markets: an overview.* (NBER Working Paper No. 16428). The National Bureau of Economic Research. https://www.nber.org/papers/w16428.pdf

Rohrer, J. M. (2018). Thinking Clearly About Correlations and Causation: Graphical Causal Models for Observational Data. *Advances in Methods and Practices in Psychological Science, 1* (1), 27–42. https://doi.org/10.1177/2515245917745629

Sahay, R., Čihák, M., N'Diaye, P., Barajas, A., Ayala Pena, D., Bi, R., ...Yousefi, S.R. (2015b) *Rethinking Financial Deepening: Stability and Growth in Emerging Markets*. (No. 2015/08). International Monetary Fund. https://www.imf.org/external/pubs/ft/sdn/2015/sdn1508.pdf

Sahay, R., Čihák, M., N'Diaye, P., Barajas, A., Mitra, S., Kyobe, A., ..., & Yousefi, S.R. (2015a). *Financial inclusion: can it meet multiple macroeconomic goals?* (No. 2015/17). International Monetary Fund. https://www.imf.org/external/pubs/ft/sdn/2015/sdn1517.pdf

Saha, M. and Dutta, K.D. (2022). Do macroprudential regulations condition the role of financial inclusion for ensuring financial stability? Cross-country perspective, *International Journal of Emerging Markets*, Vol. ahead-of-print No. ahead-of-print. https://doi.org/10.1108/IJOEM-08-2021-1232

Shim, I., Bogdanova, B., Shek, J., & Subelyte, A. (2013). Database for policy actions on housing markets. *BIS Quarterly Review September*, 83–95. https://www.bis.org/publ/qtrpdf/r_qt1309i.pdf

World Bank. (2019a). The Worldwide Governance Indicator Dataset [Data file]. https://info.worldbank.org/governance/wgi/

World Bank. (2019b). The World Development Indicators (WDI) [Data file]. https://databank.worldbank.org/source/world-development-indicators

World Bank. (2019c). The Global Financial Development Database (GFDD) [Data file]. https://datacatalog.worldbank.org/dataset/global-financial-development

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APPENDIX

Table A1 Classification results of logit and probit estimated models

Models	Correctly classified	Sensitivity	False negative rate	Specificity	False positive rate
Probit	93.24%	4.85%	6.06%	99.93%	35.71%
Logit	93.31%	2.91%	5.94%	99.86%	31.25%

Table A2: Robustness test (unwinsorised data): joint impact of financial inclusion and macroprudential policy

Variables	(1) Probit	(2) Logit	(3) Marginal effect (Probit)	(4) Marginal effect (Logit)
FOI_{t-1}	-18.734***	-41.818***	-18.734***	-41.818***
	(6.829)	(14.530)	(6.829)	(14.530)
FOI^2_{t-1}	36.137*	77.424*	36.137*	77.424*
	(18.949)	(39.988)	(18.949)	(39.988)
MPI_{t-1}	-0.312	-2.082	-0.312	-2.082
	(2.638)	(5.744)	(2.638)	(5.744)
MPI^2_{t-1}	-6.804***	-11.858***	-6.804***	-11.858***
	(1.713)	(4.732)	(1.713)	(4.732)
$FOI_{t-1} \times MPI_{t-1}$	-12.055***	-19.865***	-12.055***	-19.865***
	(4.287)	(9.758)	(4.287)	(9.758)
CON_{t-1}	-1.479***	-3.172***	-1.479***	-3.172***
	(0.526)	(1.146)	(0.526)	(1.146)
FF_{t-1}	-0.111	-0.550	-0.111	-0.550
	(0.879)	(1.788)	(0.879)	(1.788)
TG_{t-1}	1.516**	3.375**	1.516**	3.375**
	(0.695)	(1.433)	(0.695)	(1.433)
GDP_{t-1}	-0.435	-0.343	-0.435	-0.343
	(1.778)	(3.621)	(1.778)	(3.621)
$gGDPC_{t-1}$	-0.934	-1.981*	-0.934	-1.981*
	(0.607)	(1.183)	(0.607)	(1.183)
GOV_{t-1}	-0.922	-2.016	-0.922	-2.016
	(0.965)	(2.025)	(0.965)	(2.025)

CPI_{t-1}	-0.523	-0.246	-0.523	-0.246
	(0.839)	(1.659)	(0.839)	(1.659)
FID_{t-1}	-2.576**	-5.708**	-2.576**	-5.708**
	(1.150)	(2.393)	(1.150)	(2.393)
POP_{t-1}	1.444*	2.991*	1.444*	2.991*
	(0.789)	(1.576)	(0.789)	(1.576)
Constant	-1.369*	-2.518*	-1.369*	-2.518*
	(0.734)	(1.476)	(0.734)	(1.476)
Observations	1,494	1,494	1,494	1,494

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*MilutinJešić**

DRIVERS OF GDP GROWTH: EVIDENCE FROM SELECTED EUROPEAN COUNTRIES**

ABSTRACT: *This empirical study analyses the potential determinants of GDP growth in selected European countries using data for 19 countries from Central, Eastern, and South-Eastern Europe for the period 2014 to 2020. The influence of possible drivers of economic growth are investigated by employing dynamic panel data modelling, specifically the system GMM method. The study's findings reveal that fiscal responsibility, initial development, the inflation rate, and EU membership are the main GDP growth drivers. In addition, we control for the institutional determinants of economic growth and the role of R&D. These results provide further support for the hypothesis that macroeconomic policies conducted in a responsible and sustainable way can significantly improve countries' growth perspectives. These findings may help us to understand that the trinity of policies, institutions, and technology is a conditio sine qua non of economic growth.*

KEY WORDS: *GDP growth, GDP growth drivers, selected European countries, dynamic panel data, System GMM*

JEL CLASSIFICATION: C23, C54, E32, O47, O52

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1. INTRODUCTION

Sustainable growth is the ultimate goal of every economy. Various macroeconomic, sectoral, and microeconomic policies can contribute to achieving this aim. However, sustainable growth is not only based on policy, it is rather a trinity of policies, institutions, and technology. Although we are proponents of the importance of all three factors of growth, this paper is focused only on macroeconomic and some institutional variables that are potential growth drivers. Therefore, the key task for policymakers is to conduct pro-growth oriented policies. Although sustainable growth is usually long-run oriented, we will only analyse short-term GDP growth patterns in selected economies.

The empirical determination of GDP growth determinants is one of the most interesting topics for research in macroeconomics. There is some consensus that initial conditions, macroeconomic stability, and structural policies are key GDP growth determinants and are specific to countries at different levels of development. An additional factor can be membership in economic unions since free riding effects can be exploited, but negative spillover effects are also a real threat to countries that belong to the specific union.

Our motivation for this study lies in our desire to investigate the main determinants of GDP growth in selected European countries in Central and Eastern Europe. Some of them are EU members; however, some of them are only candidates for EU membership, or do not even have that status. The main research question is: what are the important determinants of GDP growth in selected European economies, primarily from the policy side? Of special interest in this study is the impact of macroeconomic policy on GDP growth. Although some policy instruments do not have a direct effect on growth, indirect effects can be significant. In addition, the study seeks to reveal whether institutions have a significant role in shaping growth patterns in the observed period.

This study is intended to fill gaps identified in the literature. First, the sample of countries observed is unique, on the one hand. On the other hand, the chosen countries have many common characteristics that prompted us to include them in the sample. Second, state-of-the-art econometric techniques are used, which assures us of the robustness of the results. Third, important policy implications are derived on the basis of the findings.

The rest of the paper is organised as follows. Section 2 provides some brief insights into the findings of empirical papers dealing with similar topics. Section 3 is dedicated to the data in general, variables used in the analysis, and brief stylised facts. Section 4 presents the methodology used in this study. Section 5 provides the results of the study and Section 6 some brief policy implications. Section 7 concludes.

2. LITERATURE REVIEW

To get a broader picture of the literature dealing with GDP growth and to provide a brief literature review, we employed VOSviewer (van Eck & Waltman, 2014; Waltman et al., 2010), which is open-source software for creating bibliometric maps[1]. We identified 3 clusters of words that build one network and gave them the names that, in our opinion, best describe the specific cluster. The clusters and the most representative ingredient words are: 1. *Macroeconomic determinants* (macroeconomic factors, panel data, emerging markets, public debt, inflation); 2. *Sustainable economy* (CO_2 emissions, energy efficiency, renewable energy, environmental Kuznets curve); 3. *Institutions and human capital* (education, convergence, income inequality, European Union, institutions). Another important observation based on overlay visualisation is that authors have focused on topics within the second cluster in recent years, while the other two clusters have been overshadowed. Of course, the clusters are strongly connected and sometimes the same paper can be classified in various clusters based on the words in focus.

Due to the baseline research question and our results, our paper can be classified into the first and third cluster. Therefore, in the following brief literature review we will focus on empirical papers that analyse GDP growth patterns and especially the role of macroeconomic factors of GDP growth.

One strand of the literature deals with broad macroeconomic determinants of GDP growth. Barro (1999) investigates factors of economic growth in a panel of 100 countries for the period 1960–1995, finding that investment share, terms of trade, years of schooling, the rule of law index, the democracy index, and

1 The details on this particular analysis in VOSviewer and Illustration that represents mapping of the words and identified clusters can be found in supplementary material here: https://doi.org/10.7910/DVN/WRSJNJ

international openness positively influence economic growth, while government consumption, the total fertility rate, and inflation negatively impact it. Bayraktar (2006) analyse the relationship between growth rates and various macroeconomic indicators in Turkey for the period 1968–1998, and conclude that investments, human capital development, and inflation are determinants of Turkish economic growth. Trpkova and Tashevska (2011) find that CPI, the current account, the exchange rate, general government balance, general government expenditure, and population are some of the main factors of GDP growth in seven Southeast European countries for the 1995–2007 period. Checherita-Westphal and Rother (2012) examine the role of government debt level and economic growth in twelve selected euro area countries for the 1970–2008 period. Their results indicate that government balance, private savings, and trade openness are positively related to economic growth, while population growth and real interest rates are negatively related. In addition, government debt positively affects economic growth. However, the squared variable of government debt negatively impacts growth, suggesting non-linear behaviour of the relationship and the presence of threshold effects, i.e., above a certain government debt level, negative effects become dominant. Prochniak (2011) attempts to find the determinants of growth in ten CEE economies during the 1993–2009 period by employing OLS estimation. He argues that the investment rate, human capital development, financial sector development, high services share in GDP, high share of working age population, development of information, communication and technology, high private sector share in GDP, economic freedom, and progress in market and structural reforms positively affect economic growth, while budget deficits, public debt, interest rates, and inflation negatively affect it. Josifidis et al. (2012) investigate the causes of the heterogeneity in growth rates in the Western Balkan and emerging European economies in the period 1997–2009 using dynamic panel data models. They find that macroeconomic stabilisation and reforms are significant determinants of GDP growth, but that foreign direct investments and economic integrations are the key determinants. Schneider and Wagner (2012) use the adaptive Lasso estimator to determine factors of economic growth for a regional dataset for the European Union covering the 255 NUTS2 regions in the 27 member states over the period 1995–2005. They find that initial GDP per capita, human capital, and structural labour market characteristics are important for GDP growth. Fetahi-Vehapi et al. (2015) analyse determinants of economic growth in ten South-Eastern European countries during the period 1996–2012 using a fixed effects panel regression estimation method and find that trade openness, GDP per capita, human capital development, gross fixed capital formation, and foreign direct investment positively impact economic growth, while population negatively impacts it. Simionescu et al. (2017) study drivers that might influence GDP growth in the Czech Republic, Slovakia, Hungary, Poland, and Romania for the period 2003–2016. Among other results, expenditure on education impacted economic growth only in the Czech Republic, while expenditure on R&D had positive effects in Romania, Hungary, and the Czech Republic. In an empirical analysis of the determinants of inclusive growth between 1980 and 2013 for a sample of 78 countries, Jalles and de Mello (2019) find that important drivers of inclusive growth are human capital accumulation, the redistributive potential of tax-benefit systems, increases in multifactor productivity and labour force participation, trade openness, and some institutional factors. D'Andrea (2022), investigating the determinants of growth in 19 European countries from 2002 to 2019, concludes that there are several robust determinants of growth, such as the initial level of GDP per capita, savings, the share of manufacturing in GDP, demography, public accounts, wage and labour contract regulation, and fixed capital accumulation.

The second strand of the literature is more specialised. Bleaney et al. (2001) analyse the connection between fiscal policies and economic growth in 22 developed countries for the 1970–1995 period. They argue that productive government expenditure has a significant positive influence, while distortionary fiscal policies have a negative influence on the long-run economic growth rate. Bittencourt (2012) investigate the relationship between inflation and economic growth in Latin America for the period between 1970 and 2007. Their results suggest that inflation has a detrimental effect on growth in the region. Baum et al. (2013), analysing the relationship between public debt and economic growth in twelve euro area countries for the period 1990–2010, find nonlinearity and show that low levels of public debt have a positive impact on GDP growth, but this converges to zero with rising debt. Beyond a threshold of around 95%, additional debt has a negative impact on economic activity. Dreyer and Schmid (2017), studying the growth effects of EU and eurozone memberships on data for the first 15 years of the euro – from 1999 to 2013, argue that a positive impact of EU membership on economic growth is present, although being part of the eurozone has no impact on growth, except the negative effect during the financial crisis.

Ješić and Jakšić (2020) theoretically and empirically establish a relationship between institutional features and R&D in the business enterprise sector and, consequently, on sustainable economic growth in the data for eight European countries for the period 2007–2017. Arsić et al. (2021) examine the response of economic growth to public debt uncertainty in ten emerging European economies between 2000 and 2015. Their results indicate the negative effect that public debt uncertainty has on GDP growth in emerging European economies. In addition, such an impact was amplified during the 2008 crisis episode. Kassouri et al. (2021) analyse the nexus between debt and growth in a sample of 62 emerging and developing countries from 2000 to 2018, with their results indicating the presence of an inverted U-shaped relationship between debt and growth. By employing a dynamic panel, they find that public debt harms growth when the indebtedness level exceeds the estimated threshold of 50.19% and 25.09% of GDP for the upper-middle income and low income subsamples, respectively. In addition, these negative effects are of higher intensity in low income countries. Law et al. (2021) employ a dynamic panel threshold technique to provide new evidence on the threshold value of the ratio of public debt to the gross domestic product in 71 developing countries from 1984 to 2015. Debt has a negative influence on economic growth at a high level of public debt (threshold value is 51.65%) but an insignificant effect at a low level of public debt.

The third strand of the literature is mainly focused on the role of institutional development and its impact on economic growth. Iqbal and Daly (2014) argue that weak institutions, especially characterised by a low level of the rule of law and inadequate political and public policies, have a negative influence on GDP growth. Acemoglu and Robinson (2010) find that divergences between countries relating to GDP growth are primarily caused by institutions. Shapkova and Disoska (2017), investigating the impact of institutions on economic growth in transition economies in Central and Eastern Europe and the Western Balkans in the period from 2000 to 2016, find that a positive relationship between economic growth and the rule of law, control of corruption, regulatory quality, and voice and accountability is visible. Radulović (2020) shows that there is a long-term and a short-term relationship between the quality of institutions (six dimensions of governance) and economic growth in ten observed SEE countries (five EU members and five non-EU members) from 1996 to 2017.

3. DATA AND STYLISED FACTS

In this study, the drivers of GDP growth are analysed using a sample of the following countries: Belarus, Bosnia and Herzegovina, Bulgaria, Croatia, Cyprus, the Czech Republic, Estonia, Greece, Hungary, Latvia, Lithuania, Poland, Romania, the Russian Federation, Serbia, Slovakia, Slovenia, Türkiye, and Ukraine. Regarding the sample of countries used, it is, on the one hand, heterogeneous. According to World Bank classification, some of these countries are high income countries, some are upper-middle income countries, and one is a lower-middle income country. On the other hand, these countries have many common characteristics. First, their starting position before the transition process was similar. Second, the majority of the sample countries were part of the former socialist block. Third, all the countries belong to the wider area of Central, Eastern and South-Eastern Europe. Fourth, the determinants of growth in these countries are certainly more similar than those of Western economies. Consequently, the non-EU countries can observe important policy implications from the GDP growth patterns of the EU countries. The time span ranges from 2014 to 2020.

The data sources used in the analysis are the World Bank's World Development Indicators (WDI) database (2022a), the IMF's World Economic Outlook (WEO) database (2022) and the World Bank's World Governance Indicators (WGI) database (2022b). The sources and description of the data are given in the Table 1.

Table 1. Sources and data description

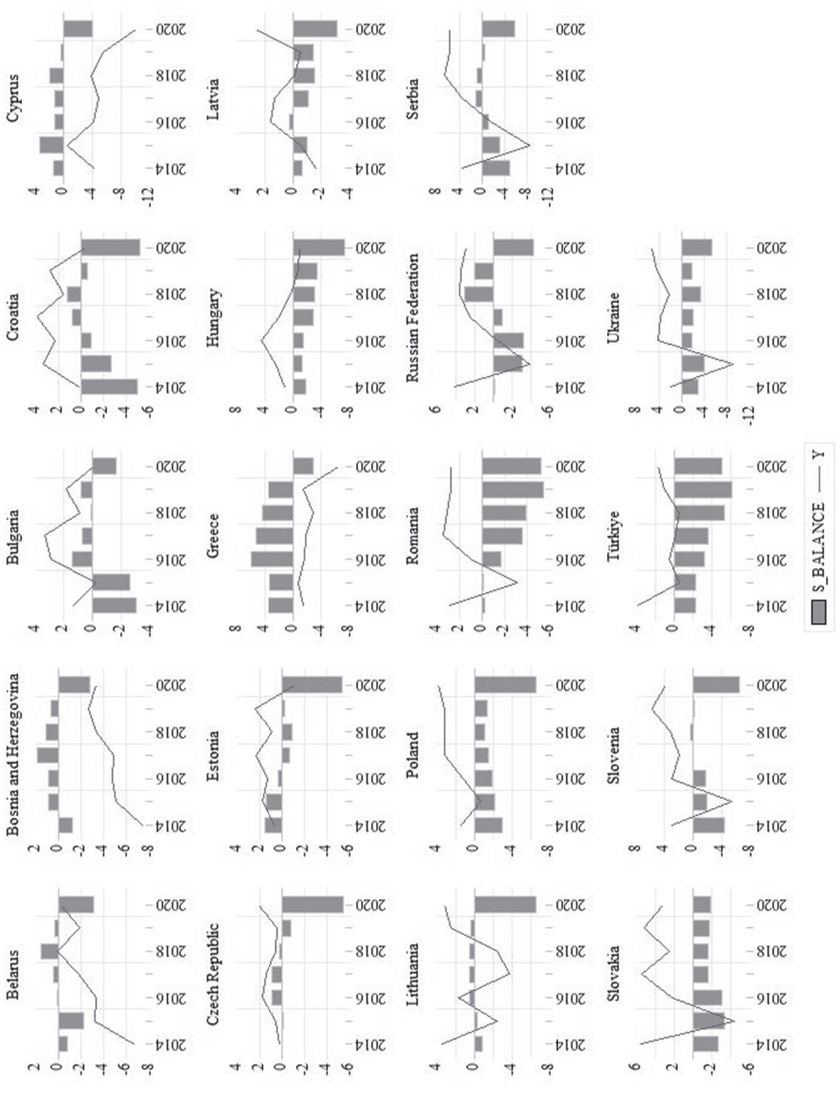
Variables	Description	Source
Y	GDP growth	World Bank, WDI database
DEBT	Public debt in % of GDP.	IMF, WEO database
S_BALANCE	Structural budget balance in percentage of potential GDP.	IMF, WEO database
GDPPC	Logarithm of GDP per capita	World Bank, WDI database
CPI	CPI measured inflation at the end of the period	World Bank, WDI database
R&D	Research and development expenditure (% of GDP)	World Bank, WDI database
RULE_LAW	Rule of Law. Ranges from approximately -2.5 (weak) to 2.5 (strong) governance performance.	World Bank, WGI database
CRISIS	Takes value of 1 in 2020, and 0 otherwise.	
EU	Takes value of 1 if the country is member of the European Union, and 0 otherwise.	

Our analysis is dedicated to the investigation of possible determinants of GDP growth (*Y*), which is our model's dependent variable. Due to the expected dynamics and persistence of adjustment, the regressors in our analysis will be lagged values of GDP growth and additional various covariates, usually macroeconomic ones. In the rest of this section, we will briefly propose some hypotheses on possible covariates influencing GDP growth and provide some stylised facts.

Fiscal policy can have beneficial or detrimental effects on GDP growth, depending on the way it is conducted. The measurement of public debt (*DEBT*) sustainability is an open question in fiscal policy studies, and there is no consensus on what levels of public debt are unsustainable and what other prerequisites should be observed in the analysis of public debt sustainability. However, there is some consensus that low- and middle-income countries are constrained in this sense at lower levels of public debt. Additionally, notwithstanding some exceptions, the dominant stance in the literature is that the negative influence of public debt on economic growth prevails. Ultimately, investors will have the final say on this issue through the interest rate adjustment mechanism, and consequently on GDP growth.

One more variable from the fiscal area that can be of great importance in analysing GDP growth patterns is the structural budget balance (*S_BALANCE*). We find it very important as this variable is used to measure fiscal (ir)responsibility, which consequently alters GDP growth through various channels of the transmission mechanism. The structural budget balance is a direct measure of the discretionary policy of the fiscal policymaker. When we exclude the cyclical component and other one-off measures, we can observe the actual underlying behavior of the government and the size of its impact on GDP growth. The concept of the structural budget balance can be of great help in concluding whether the fiscal policymaker has behaved in a procyclical or countercyclical way. Although the overall budget balance can be used to determine procyclicality / countercyclicality, this approach does not give a clear picture and, consequently, may be misguiding in evaluating the policy stance. Since the cyclical component is part of the overall budget deficit and not under the direct control of policymakers, it can be hard to distinguish whether the policy orientation is skewed to procyclicality or countercyclicality. We analysed the data for the observed countries and present them in the Figure 1. The bars represent the structural fiscal balance, i.e., positive bars mean structural budget surpluses, while negative bars mean structural budget deficits. Matching them with the GDP growth data, represented by the lines, enables us to derive some conclusions based on these stylised data. In some of the observed countries, policymakers behaved procyclically in many periods, injecting fiscal stimuli in periods of GDP growth and tightening fiscal policy in periods of recessions. However, we believe in the positive impact of fiscal responsibility proxied by the structural budget balance on GDP growth, and therefore expect a positive sign of the impact.

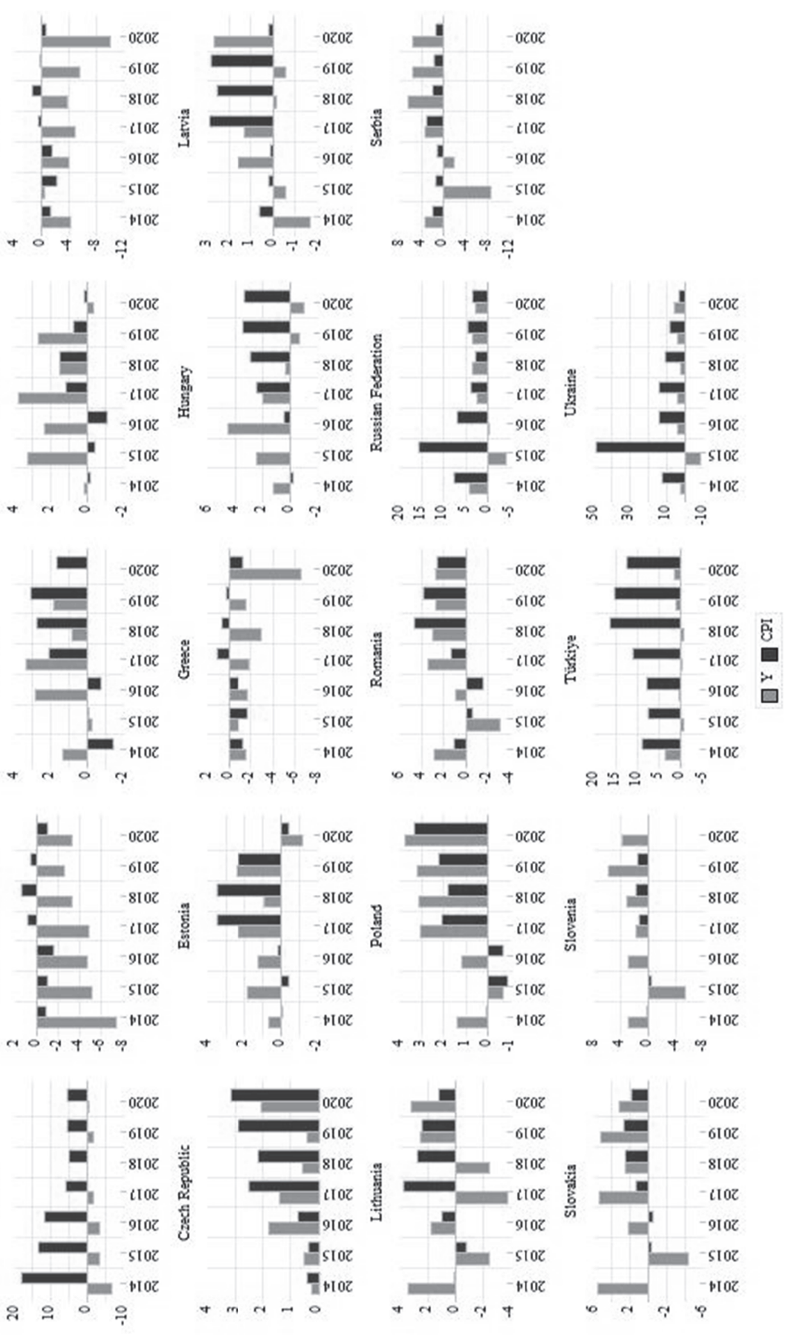
Figure 1. Fiscal (pro/counter)cyclicality and GDP growth



One important determinant of GDP growth recognised by empirical studies relates to the initial conditions. In order to investigate this potential determinant, we include the logarithm of GDP per capita (*GDPPC*). There is no consensus in the literature about the expected sign of the influence, since one strand of literature argues that in low- and middle-income countries more scope is available for growth, while the second strand argues the opposite, relating it to the level of development, which can be stimulating to growth. Therefore, the expected sign of the influence is unclear. Considering our group of observed countries and the findings of similar empirical studies, we are more willing to expect a positive sign of the respective coefficient.

Many empirical studies include the inflation rate as a covariate in the analysis on the basis of theoretical foundations, and we decided to do the same. The inflation rate (*CPI*) can be a significant driver of GDP growth tendencies. However, the expected sign of its influence is not clear-cut and depends on the specific economic conditions of the respective country. On the one hand, low inflation and low inflation volatility can be supportive to economic growth under certain conditions. In this situation, the central bank does not fear a deflation trap, the monetary policy instrument set is wider, agents interact in a stable economic environment, the government tax base is stable, and investors are optimistic and invest more, which has an overall positive impact on GDP growth. On the other hand, high inflation or high volatility of inflation are destimulative to GDP growth. In this situation, risks increase and agents behave in a way that is not conducive to GDP growth. Although some countries in our panel of countries have a very high level of inflation, the majority of them performed very well on this criterion in the observed period, which can be seen in Figure 2. However, as we already emphasised, even relatively low inflation and low inflation volatility sometimes can be a burden to economic growth.

Figure 2. *Y–CPI* relationship in the observed countries.



Research and development expenditure as % of GDP (*R&D*) is a very important determinant of economic growth, and this variable can be found in various growth theories. It can be good proxy for technology as a growth driver. Even though this variable is not always found in macro-oriented empirical papers, we still want to control for it due to theoretical considerations.

Rule of law (*RULE_LAW*) is a variable that can affect GDP growth, and we can classify it as an institutional determinant of growth. We extract the data from the World Governance Indicators database of the World Bank (2022b). It ranges from approximately -2.5 (weak) to 2.5 (strong) governance performance. We expect a more developed institutional framework to positively influence GDP growth.

The variable *CRISIS* is a dummy variable that takes the value of 1 in 2020, and 0 otherwise. The reason for the inclusion of this variable is the obvious shock that hit all countries during the initial phase of the pandemic.

Finally, we control for European Union membership (*EU*) as a possible driver of growth and for possible interaction between this variable and other covariates in the model specifications. In our sample, 13 out of 19 countries are members of the EU. Given the proven advantages of being part of the EU in terms of GDP growth, we expect a positive sign for this particular coefficient.

As mentioned, although the panel of countries is heterogeneous, there are many common characteristics of these countries. The following Table 2 provides a summary of the descriptive statistics regarding the variables used.

Table 2. Summary descriptive statistics

Variable	Mean	Std. Dev.	Min	Max
Y	1.9492	3.3408	-10.0789	8.1965
DEBT	56.5851	38.9488	8.2040	212.4490
S_BALANCE	-1.3100	2.5477	-7.3340	5.8880
GDPPC	4.0900	0.2508	3.3273	4.4686
CPI	2.9862	5.7795	-2.0970	48.6999
R&D	0.9764	0.4793	0.1926	2.3655
RULE_LAW	0.2904	0.6612	-1.0537	1.3728

Source: Author's calculation

4. METHODOLOGY

We employed a dynamic panel data estimation method, although modelling in a panel data framework can be done using various techniques, such as a pooled OLS method, fixed and random effects estimation, and instrumental variables methods. We opted for this method because of the high probability of latency in the dependent variable path in our specification. That has important consequences, since in these situations estimation of the model using the standard panel data techniques results in biased and inconsistent estimations. Many empirical studies use a fixed effects estimation (FE), but this can cause bias, especially if the number of periods is low (Nickell, 1981; Kiviet, 1995). Therefore, we opted for standard dynamic panel data modelling. However, there are different estimators that can be used in this field, such as the Arellano and Bond (1991), Arellano and Bover (1995), and the Blundell and Bond (1998) methods. We used Blundell and Bond's one-step procedure (system GMM method).

The system GMM method accounts for endogeneity of the lagged dependent variable and is asymptotically more efficient. In addition, it reduces finite sample bias and is dominant according to this criterium in comparison with other dynamic panel data estimators (Baltagi, 2008). Soto (2009) finds that this method gives the best results in the case of small N, as is the case in our study, and that its application to small samples does not have significant repercussions for the properties of the estimator. System GMM assumes a system of equations, where lagged first differences of the dependent variable are instruments for the equations in level, whereas lagged levels of the dependent variable are used as instruments for equations in first differences (Blundell & Bond, 1998). Instruments for other endogenous variables can be used. Roodman (2009) investigated a potential problem with too many instruments and this was a real risk in our panel due to the small N. It can be solved by limiting the number of lags and collapsing the instrument matrix. We opted to use both methods to limit the number of instruments to a desirable level (where the number of instruments is lower than N).

Our model specification can be expressed in the following way:

$$Y_{it} = \rho Y_{it-1} + \alpha DEBT_{it} + \beta S\_BALANCE_{it} + \delta \boldsymbol{X}'_{it} + \mu_i + \varepsilon_{it}, \tag{1}$$

where *Y* is GDP growth, *DEBT* is public debt, *S_BALANCE* is the structural budget balance, and ***X*** is a vector of control variables, such as logarithm of GDP per capita (*GDPPC*), CPI measuring the inflation rate (*CPI*), research and development in % of GDP (*R&D*), rule of law (*RULE_LAW*), the crisis dummy variable (*CRISIS*), the EU dummy variable (*EU*), and interaction terms between particular variables, while μ_i captures unobserved country-specific effects and ε_{it} is the error term.

The Blundell and Bond estimator requires stationarity of all the variables used. We employed first generation panel unit root tests, which have advantages in small samples (Breitung, 2001). Therefore, we used the following panel unit root tests: the ADF Fisher-type test, the Hadri LM test, and the Breitung test. The results of the unit root testing are presented in Table 3. The results are mixed and although some of the tests find unit root processes of some variables, we proceed to estimation due to the results of other tests.

Table 3. Unit root tests results

	ADF Fisher		**Hadri LM**		**Breitung**	
	Statistic	p value	Statistic	p value	Statistic	p value
Y	22.3250	0.9798	1.1794	0.1191	-3.3739	0.0004
DEBT	32.0242	0.7413	4.1141	0.0000	-2.5076	0.0061
S_BALANCE	58.7735	0.0169	2.6334	0.0042	-1.1130	0.1328
GDPPC	65.5454	0.0036	6.7875	0.0000	-1.1797	0.1191
CPI	22.8344	0.9754	4.1011	0.0000	-1.3939	0.0817
R&D	85.8010	0.0000	7.1617	0.0000	1.4493	0.9264
RULE_LAW	34.9925	0.6093	3.4622	0.0003	-1.5563	0.0598

Source: Author’s calculation

The adequacy of the method used is tested by standard statistical tests in this field. The Arellano–Bond test is used to investigate the presence of second-order serial correlation of the differenced residuals. The Hansen test for overidentifying restrictions is used to investigate the instruments’ validity, i.e., their exogeneity.

5. RESULTS

The results of the baseline specifications are presented in Table 4. The estimated model specifications possess good statistical properties according to standard criteria and can be used for further statistical inference.

Generally speaking, the presented results are as expected and the initial hypotheses about the determinants of economic growth are confirmed. We identified the most important determinants of GDP growth. All specifications have some common variables that significantly influence GDP growth, and according to our results these variables come from the fiscal policy side. Initial conditions are estimated to be significant, as is the variable that captures the shock caused by the pandemic in 2020. The influence of variable(s) from the monetary policy side is also sometimes examined in the literature, which motivated us to take a similar approach in baseline specification No. 3.

It is apparent from these results that the GDP growth variable shows persistence in its path, especially in the third specification. This clearly justifies our decision to use dynamic panel data model analysis.

Public debt also negatively influences GDP growth in the observed countries, although the value of the coefficient is not high in any specification. This is not only already a well-established relation in theory, but also in empirical studies (Barro, 1999; Bleaney et al., 2001; Prochniak, 2011; Arsić et al., 2021). If we compare our results to findings of other empirical papers focusing on threshold levels beyond which public debt starts to negatively influence GDP growth, our results are contrary to the findings of Baum et al. (2013), and relatively consistent with the findings of Kassouri et al. (2021) and Law et al. (2021). Only two countries stand out in terms of the level of public debt in the observed period, Greece and Cyprus. Our results support the findings of other studies that this threshold is certainly lower for emerging market economies.

Table 4. Estimated baseline models

Dependent variable: *Y*	Model 1	Model 2	Model 3
Y (-1)	0.2111*	0.2548**	0.7941***
	(0.1269)	(0.1192)	(0.1467)
DEBT	-0.0858**	-0.0790**	-0.0901***
	(0.0336)	(0.0341)	(0.0348)
S_BALANCE	0.5471**	0.5611***	0.8487***
	(0.2420)	(0.1969)	(0.2032)
GDPPC	1.8623***	1.7426***	
	(0.4567)	(0.4788)	
CRISIS	-3.7658***	-6.3040***	
	(1.3231)	(1.5817)	
*CRISIS*S_BALANCE*		-0.5440**	
		(0.2644)	
CPI			0.3855***
			(0.1309)
EU			6.9806***
			(2.2524)
*EU*CPI*			-1.0695**
			(0.5016)
Arellano-Bond AR(2) test	-1.04 [0.2960]	-0.85 [0.3960]	-0.20 [0.8430]
Hansen test	13.75 [0.1320]	14.50 [0.1060]	16.14 [0.0960]
No. of observations	114	114	114
No. of countries	19	19	19

Notes: ***, **, * denote statistical significance at the 1%, 5%, and 10% levels, respectively. Robust standard errors are in parentheses. System GMM with robust standard errors is applied. Instruments used for the level equation are lagged first differences of *Y*, *DEBT*, *GDPPC*, and *CPI* (potentially endogenous variables). Instruments used for the first-differenced equations are lagged levels (three period) of the dependent variable, and of the potentially endogenous variables. All other variables are treated as exogenous covariates and are instrumented by themselves in the level equations. The p values for the Arellano-Bond and Hansen tests are in square brackets.
Source: Author's calculation

Another variable from the fiscal domain also contributes significantly in all our model specifications – the structural budget balance. The value of the coefficient is quite high, especially in the third specification. This is evidence in favour of the importance of the hypothesis that fiscal responsibility contributes significantly to

GDP growth and that properly conducted discretionary fiscal policy can support economic growth. In addition, these findings are consistent with the role of public debt as a GDP growth driver, jointly providing a strong rationale for emphasising these fiscal policy variables.

We found that the initial level of development proxied by GDP per capita achieved a powerful positive effect on GDP growth. This finding is consistent with the those of Fetahi-Vehapi et al. (2015) and Schneider and Wagner (2012). As has already been stated, low- and middle-income economies have a lower base and are better able to achieve higher growth rates. However, the volatility of growth rates is not as high and there are high spillover effects from developed countries. This is the reason for the positive sign of this coefficient.

As expected, dummy variables also influence GDP growth. The crisis variable shows that in 2020 many countries suffered from the effects of the pandemic. Finally, being part of the EU significantly increases GDP growth (if CPI is not too high), which is consistent with the findings of Dreyer and Schmid (2017).

In addition to these variables, we aimed to control for possible interactions between covariates. The first interaction term in model specification No. 2 indicates the significant influence of the crisis on the structural balance in the sense that for the non-crisis years the coefficient of influence of the structural balance on GDP growth is 0.5611. However, if we observe solely the crisis year, this coefficient is 0.0171 (0.5611-0.5440). Therefore, the impact of the structural balance on GDP growth depends on the effects of the crisis.

The second interaction term is also very important because it shapes the overall effect of the inflation rate on GDP growth. In this case, the effect of inflation on GDP growth is positive for non-EU countries (0.3855). However, things are quite different if we take into account this interaction term, because the overall coefficient is -0.6840 (0.3855-1.0695) for EU countries, which suggests the negative impact of inflation on GDP growth in EU countries. Due to the relatively low level of inflation in the observed period, low inflation was not harmful to economic growth in non-EU countries, in contrast to the situation of EU countries, whose growth is more sensitive to inflation, even if it is not high, due to more liberalised cross-border trade. Concerning this interaction between the EU and CPI variables, our main angle of view is from the side of the CPI variable, and the EU variable has the function of a control variable. However, we can also analyse this from the other angle and conclude that if CPI is small or even equal to zero, the EU effect is extremely powerful, but if CPI is high, the EU effect diminishes progressively to zero beyond an inflation rate of 6.53%, which can act as some threshold. In other words, the net effect of the EU variable depends on the level of inflation since the marginal effect of this variable is a linear function of the CPI level[2]. The explanation can be found in the underlying data. Namely, if CPI is low (which was the case in many EU and non-EU sample countries in the observed period), then the EU variable is powerful and being part of the EU has positive consequences for GDP growth. However, if CPI is high enough, above a certain threshold, then the EU effect is insignificant due to the data facts (inflation above this threshold was recorded only in the following non-EU countries during the observed period: Belarus, the Russian Federation, Türkiye and Ukraine).

As was stated earlier in this paper, theory suggests that potential drivers of GDP growth can be the institutional setup in a country and technology (Iqbal & Daly, 2014; Acemoglu & Robinson, 2010; Shapkova & Disoska, 2017; Radulović, 2020). Although our primary goal was to identify GDP growth drivers from the policy side, we are aware of the importance of other variables that could have an important impact on economic growth. In order to control for these, they are included in the alternative model specifications. Models 4–6 correspond to the baseline models 1–3, respectively, but now including these two potential determinants of economic growth as control variables (rule of law and research and development).

2 The marginal effect of the EU variable is equal to the coefficient of the EU variable + the coefficient of the interaction term * CPI.

Table 5. Estimated alternative models

Dependent variable: *Y*	Model 4	Model 5	Model 6
Y (-1)	0.1890	0.2312*	0.5512***
	(0.1333)	(0.1186)	(0.2149)
DEBT	-0.0868**	-0.0811**	-0.1074***
	(0.0347)	(0.0349)	(0.0370)
S_BALANCE	0.5630**	0.5793***	0.9633***
	(0.2549)	(0.2100)	(0.2037)
GDPPC	1.8182***	1.7535***	
	(0.5656)	(0.5944)	
CRISIS	-3.7415***	-6.2384***	
	(1.2755)	(1.7822)	
*CRISIS*S_BALANCE*		-0.5468*	
		(0.3099)	
CPI			0.3098***
			(0.1090)
EU			5.7550***
			(2.0486)
*EU*CPI*			-1.0537**
			(0.4201)
R&D	0.2333	0.0659	2.7165
	(1.2899)	(1.2326)	(1.6775)
RULE_LAW	0.3680	0.3345	1.2066
	(0.5125)	(0.4791)	(0.8655)
Arrelano-Bond AR(2) test	-1.03 [0.3050]	-0.86 [0.3910]	-0.35 [0.7230]
Hansen test	14.04 [0.1210]	14.18 [0.1160]	14.62 [0.1460]
No. of observations	114	114	114
No. of countries	19	19	19

Notes: ***, **, * denote statistical significance at the 1%, 5%, and 10% levels, respectively. Robust standard errors are in parentheses. System GMM with robust standard errors is applied. Instruments used for the level equation are lagged first differences of *Y*, *DEBT*, *GDPPC*, and *CPI* (potentially endogenous variables). Instruments used for the first-differenced equations are lagged levels (three period) of the dependent variable, and of the potentially endogenous variables. All other variables are treated as exogenous covariates and are instrumented by themselves in the level equations. The p values for the Arellano-Bond and Hansen tests are in square brackets.

As was the case in the baseline estimations, the estimated alternative model specifications possess good statistical properties according to standard criteria and can be used for further statistical inference. The results indicate a high level of robustness when we consider these additional variables.

The research and development and rule of law variables have the expected sign of influence but are not significant (in model 6, the p value for the significance of the R&D coefficient is 0.105). We nevertheless wanted to control for them because of their potential impact according to economic theory. The other variables have a similar impact on GDP growth as in the baseline specifications, which can be seen as a good robustness check.

6. POLICY IMPLICATIONS

Based on our results, we are able to provide some brief policy implications that may be useful to policymakers not only in the observed countries, but also in many other countries at a similar stage of development. The findings of this study complement those of earlier studies in general, and the insights gained from this study may be of assistance in bridging the identified gaps that slow down GDP growth.

The results of the study highlight the importance of fiscal responsibility for sustainable GDP growth. Fiscal policy is one of the main countercyclical tools in macroeconomic policymaking due to the strong mechanism of automatic stabilisers as well as discretionary measures. However, sometimes these discretionary measures are misused by politicians in power to achieve specific goals. Moreover, fiscal policy is not delegated to an independent authority, which makes fiscal policy useful for achieving particular goals in the sphere of political macroeconomics. Such policies usually lead to procyclical expansion, which poses a real threat to fiscal sustainability. The worsening of fiscal conditions ultimately leads to decreased GDP growth. Despite not being the focus of this study, microeconomic aspects of fiscal policy are also important determinants of GDP growth, especially the structure of public expenditures. The negative effect of public debt on GDP growth was found to be small, but significant. Although we did not explicitly analyse the possibility of a reversal in GDP growth in response to an increase in public debt, we can argue on the basis of the results that if there is some threshold beyond which negative consequences of debt become apparent, this threshold is lower for the countries in our sample than in advanced economies, especially those that are not part of the EU and do not have full access to European funds. Regarding the role of the structural budget balance in explaining GDP growth patterns, clear policy implications emerge from the results. The relatively strong impact suggests the need for responsible fiscal policy, more precisely discretionary policy measures.

Notwithstanding the relatively limited sample, this work offers valuable insights into the role of monetary policy in GDP growth. The results indicate that the effect of inflation on GDP growth is positive for non-EU countries and negative for EU countries. These findings suggest several courses of action for the policymakers in both groups of countries. Even relatively low inflation can be harmful to GDP growth and, keeping in mind the world's current situation, in which inflation has reached its highest levels in the last few decades, policymakers have to be very careful in using unconventional monetary policy measures.

It could be argued that the crisis caused by the pandemic significantly influenced GDP growth. Countries in that period used various types of policy measures directed at mitigating the negative effects of this extreme shock. Some of them used unconventional measures. Therefore, in the post-crisis period, special attention should be placed on the right approach to gradually transitioning to standard policy measures.

The findings of this study raise intriguing questions regarding the role of EU membership in economic growth. In general, it seems that being part of a strong economic union is beneficial for countries at this stage of development. The European Union countries in our panel performed better in terms of economic growth. The clear implication is that countries that are not yet members of the Union have to converge more quickly to the established high standards. Even if they are not formally part of the Union, they have to adopt all best practices and be ready for formal membership. This will increase their convergence to the Union's standards in many spheres and also lead to sustainability of GDP growth. Although countries that are part of the EU can exploit some benefits and act as free riders, especially in the short run, sustainable growth is a more challenging task and requires individual efforts of the policymakers in the respective countries. This “EU effect” in the observed sample operates through the various channels, e.g. exploitation of leading EU countries credibility, economic union benefits, weaker barriers to trade, geographical and geopolitical closeness to advanced western countries, among other things. It is hoped that the countries of peripheral South East Europe will become members of the EU and this will certainly have a powerful effect on their economies. The evidence supporting this can be found in the history of EU enlargement. During the EU accession process, countries will have the scope to improve their policies and institutions to be more pro-growth oriented.

There are still many unanswered questions about the role of R&D as a GDP growth determinant in this study, but we are sure that the role is not negligible. Theory suggests that R&D is one of the main drivers of GDP growth and, bearing in mind that there are significant differences between the observed countries, policies that are pro-growth oriented have to support R&D, recognising its role in achieving the ultimate goal of economic policy. There is, therefore, a definite need for a significant step forward in these countries in order to be more competitive in a world market. Looking at the data for these countries reveals that there are huge differences between the level of R&D in percentage of GDP, e.g. in Bosnia and Herzegovina this level is around 9.8 times lower than in Slovenia on average. It is also noticeable that EU countries have higher levels of R&D investment than non-EU countries.

Finally, it is likely that connections exist between institutional development and GDP growth. Policies cannot provide full support to economic growth in the absence of institutional support. These countries are heterogeneous with respect to this criterion. Greater efforts are needed to ensure the development of inclusive institutions that promote GDP growth. Remaining in the same place is actually a step backwards. Weak institutions are an obstacle to growth since agents have to operate in a world of uncertainty. One important warning concerning institutional development is that establishing good institutions is not enough if they have no power to act. Policymakers have to decrease the gap between *de jure* and *de facto* power of good institutions.

Taken together, these results suggest that the trinity of policies, technology, and institutions is the most important precondition for growth and that ensuring responsible policies and support of technological and institutional development should be a priority for the policymakers in all the observed countries. Achieving this is a guarantee for sustainable growth.

7. CONCLUSIONS

In this investigation, the aim was to assess the determinants of GDP growth for 19 European countries in the 2014–2020 period. The panel of countries consisted of Belarus, Bosnia and Herzegovina, Bulgaria, Croatia, Cyprus, the Czech Republic, Estonia, Greece, Hungary, Latvia, Lithuania, Poland, Romania, the Russian Federation, Serbia, Slovakia, Slovenia, Türkiye, and Ukraine. The standard dynamic panel data method was employed to conduct the analysis. More specifically, we used the system GMM method, which has numerous advantages over other methods of estimation. Moreover, this method is especially appropriate when there is a persistence in the dynamics of the dependent variable, in our case GDP growth.

This study has identified the main determinants of GDP growth in the observed countries. Fiscal policy responsibility is a *conditio sine qua non* of economic growth. This hypothesis has been confirmed with the inclusion of two variables from the fiscal policy side. Public debt significantly negatively affects GDP growth. On the other hand, the structural budget balance significantly positively influences economic growth.

Overall, this study strengthens the notion that inflation, initial conditions, economic union membership, and other macroeconomic factors have an important role in driving GDP growth. In addition to these variables, we controlled for other possible drivers identified in economic theory in order to make our estimations robust, although our primary goal was to identify GDP growth drivers from the policy side. These covariates relate to the institutional framework and technology, and we are aware of their importance as drivers of economic growth.

This analysis has provided a deeper insight into the three pillars of economic growth: policies, institutions, and technology. These insights lead to significant policy implications. The observed countries in the sample have, on the one hand, much in common. On the other hand, many differences exist between them. Specific identified drivers of growth show high volatility between these countries.

The level of inflation in some countries was relatively high, even in the period of low inflation globally. The GDP per capita is also a factor that differentiates these countries as some of them are high income countries, some are upper-middle income countries, and one is a lower-middle income country, according to World Bank classification. The level of R&D investment and institutional development measured by the rule of law variable indicate that policymakers do not attach to much attention to reducing these divergences. Dealing with all these identified gaps should be the primary task of policymakers to ensure inclusive and sustainable growth.

REFERENCES

Acemoglu, D., & Robinson, J. (2010). The role of institutions in growth and development. *Review of Economics and Institutions*, *1*(2), 1–33. https://doi.org/10.5202/rei.v1i2.14

Arellano, M., & Bond, S. (1991). Some tests of specification for panel data: Monte Carlo evidence and an application to employment equations. *The Review of Economic Studies*, *58*(2), 277. https://doi.org/10.2307/2297968

Arellano, M., & Bover, O. (1995). Another look at the instrumental variable estimation of error-components models. *Journal of Econometrics*, *68*(1), 29–51. https://doi.org/10.1016/0304-4076(94)01642-D

Arsić, M., Mladenović, Z., & Nojković, A. (2021). Debt uncertainty and economic growth in emerging European economies: Some empirical evidence. *Emerging Markets Finance and Trade*, *57*(12), 3565–3585. https://doi.org/10.1080/1540496X.2019.1700364

Baltagi, B. (2008). *Econometric Analysis of Panel Data*. John Wiley & Sons.

Barro, R.J. (1999). Determinants of Economic Growth: Implications of the Global Evidence for Chile. *Cuadernos de Economía*, *36*(107), 443–478.

Baum, A., Checherita-Westphal, C., & Rother, P. (2013). Debt and growth: New evidence for the euro area. *Journal of International Money and Finance*, *32*, 809–821. https://doi.org/10.1016/j.jimonfin.2012.07.004

Bayraktar, B. (2006). Investigation on sources of growth for Turkey. *Canadian Journal of Development Studies / Revue canadienne d'études du développement*, *27*(1), 25–38. https://doi.org/10.1080/02255189.2006.9669118

Bittencourt, M. (2012). Inflation and economic growth in Latin America: Some panel time-series evidence. *Economic Modelling*, *29*(2), 333–340. https://doi.org/10.1016/j.econmod.2011.10.018

Bleaney, M., Gemmell, N., & Kneller, R. (2001). Testing the endogenous growth model: Public expenditure, taxation, and growth over the long run. *Canadian Journal of Economics/Revue canadienne d'économique*, *34*(1), 36–57. https://doi.org/10.1111/0008-4085.00061

Blundell, R., & Bond, S. (1998). Initial conditions and moment restrictions in dynamic panel data models. *Journal of Econometrics*, *87*(1), 115–143. https://doi.org/10.1016/S0304-4076(98)00009-8

Breitung, J. (2001). The local power of some unit root tests for panel data. In B.H. Baltagi, T.B. Fomby, & C.R. Hill (Eds.). *Nonstationary Panels, Panel Cointegration and Dynamic Panels, Advances in Econometrics,* JAI, 161–177. https://doi.org/10.1016/S0731-9053(00)15006-6

Checherita-Westphal, C., & Rother, P. (2012). The impact of high government debt on economic growth and its channels: An empirical investigation for the euro area. *European Economic Review*, *56*(7), 1392–1405. https://doi.org/10.1016/j.euroecorev.2012.06.007

D'Andrea, S. (2022). Are there any robust determinants of growth in Europe? A Bayesian Model Averaging approach. *International Economics*, *171*, 143–173. https://doi.org/10.1016/j.inteco.2022.06.001

Dreyer, J. K., & Schmid, P. A. (2017). Growth effects of EU and EZ memberships: Empirical findings from the first 15 years of the Euro. *Economic Modelling*, *67*, 45–54. https://doi.org/10.1016/j.econmod.2016.09.007

Fetahi-Vehapi, M., Sadiku, L., & Petkovski, M. (2015). Empirical analysis of the effects of trade openness on economic growth: An evidence for south East European countries. *Procedia Economics and Finance*, *19*, 17–26. https://doi.org/10.1016/S2212-5671(15)00004-0

IMF. (2022). World Economic Outlook database. https://www.imf.org/en/Publications/SPROLLS/world-economic-outlook-databases#sort=%40imfdate%20descending

Iqbal, N., & Daly, V. (2014). Rent-seeking opportunities and economic growth in transitional economies. *Economic Modelling*, *37*, 16–22. https://doi.org/10.1016/j.econmod.2013.10.025

Jalles, JT., & de Mello, L. (2019). Cross-country evidence on the determinants of inclusive growth episodes. *Review of Development Economics, 23*, 1818–1839. https://doi.org/10.1111/rode.12605

Ješić, M., & Jakšić, M. (2020). The Impact of Institutional Features on R&D in Business Enterprise Sector and Sustainable Growth. *Journal of Central Banking Theory and Practice*, *9*(3) 61–76. https://doi.org/10.2478/jcbtp-2020-0036

Josifidis, K., Dragutinović-Mitrović, R., & Ivančev, O. (2012). Heterogeneity of growth in the west Balkans and emerging Europe: A dynamic panel data model approach. *Panoeconomicus*, *59*(2), 157–183. https://doi.org/10.2298/PAN1202157J

Kassouri, Y., Altıntaş, H., Alancioğlu, E., & Kacou, K. Y. T. (2021). New insights on the debt-growth nexus: A combination of the interactive fixed effects and panel threshold approach. *International Economics*, *168*, 40–55. https://doi.org/10.1016/j.inteco.2021.08.001

Kiviet, J. F. (1995). On bias, inconsistency, and efficiency of various estimators in dynamic panel data models. *Journal of Econometrics*, *68*(1), 53–78. https://doi.org/10.1016/0304-4076(94)01643-E

Law, S. H., Ng, C. H., Kutan, A. M., & Law, Z. K. (2021). Public debt and economic growth in developing countries: Nonlinearity and threshold analysis. *Economic Modelling*, *98*, 26–40. https://doi.org/10.1016/j.econmod.2021.02.004

Nickell, S. (1981). Biases in dynamic models with fixed effects. *Econometrica*, *49*(6), 1417. https://doi.org/10.2307/1911408

Prochniak, M. (2011). Determinants of economic growth in central and Eastern Europe: The global crisis perspective. *Post-Communist Economies*, *23*(4), 449–468. https://doi.org/10.1080/14631377.2011.622566

Radulović, M. (2020). The impact of institutional quality on economic growth: A comparative analysis of the EU and non-EU countries of Southeast Europe. Economic Annals, *65*(225), 163–181. https://doi.org/10.2298/EKA2025163R

Roodman, D. (2009). A note on the theme of too many instruments. *Oxford Bulletin of Economics and Statistics*, *71*(1), 135–158. https://doi.org/10.1111/j.1468-0084.2008.00542.x

Schneider, U. & Wagner, M. (2012). Catching Growth Determinants with the Adaptive Lasso. *German Economic Review*, *13*(1), 71-85. https://doi.org/10.1111/j.1468-0475.2011.00541.x

Shapkova, K. K., & Disoska, E. M. (2017). Influence of trade and institutions on economic growth in transitional economies: evidences from countries from Central and Eastern Europe and Western Balkans. *Economic Analysis*, *50*(3–4), 32–42.

Simionescu M., Lazányi K., Sopková G., Dobeš K., & Balcerzak A. P. (2017). Determinants of economic growth in V4 countries and Romania. *Journal of Competitiveness*, *9*(1), 103–116. https://doi.org/10.7441/joc.2017.01.07

Soto, M. (2009). *System GMM Estimation with A Small Sample*. UFAE and IAE Working Paper 78009.

Trpkova, M., & Tashevska, B. (2011). Determinants of economic growth in South - East Europe: A panel data approach. *Perspectives of Innovations, Economics and Business (PIEB)*, *07*(1), 12–15. https://doi.org/10.22004/ag.econ.127028

van Eck, N. J., & Waltman, L. (2014). Visualizing Bibliometric Networks. In Y. Ding, R. Rousseau, & D. Wolfram (Eds.). *Measuring scholarly impact: Methods and practice* (pp. 285–320). Springer International Publishing. https://doi.org/10.1007/978-3-319-10377-8_13

Waltman, L., van Eck, N. J., & Noyons, E. C. M. (2010). A unified approach to mapping and clustering of bibliometric networks. *Journal of Informetrics*, *4*(4), 629–635. https://doi.org/10.1016/j.joi.2010.07.002

World Bank, (2022a). WDI database. https://databank.worldbank.org/source/world-development-indicators

World Bank, (2022b). WGI database. https://info.worldbank.org/governance/wgi/

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RICARDIAN EQUIVALENCE OR TWIN DEFICITS HYPOTHESIS? EVIDENCE FROM SERBIA

ABSTRACT: *This study investigates empirically whether higher budget deficits worsen the current account balance in Serbia, which is in line with the twin deficit hypothesis.This prediction is very different from Ricardian equivalence theory, which implies that a decrease in public savings is always anticipated via increased private savings. Hence, budget deficits have no impact on the current account balance. Based on quarterly data for the period between 2005 and 2020 and using a multivariate vector autoregression (VAR) model and a short-run structural VAR model, this paper confirms the twin deficit hypothesis in Serbia. More precisely, a 1 percentage point increase in the budget deficit (as a percentage of GDP) generates a 0.31 percentage point increase in the current account deficit (as a percentage of GDP). Moreover, the result is also confirmed using alternative estimation techniques (GMM and OLS method). According to these results, macroeconomic policymakers in Serbia should resort to policies that encourage fiscal consolidation to rectify or at least mitigate deterioration of the current account balance.*

KEY WORDS: *twin deficits hypothesis, Ricardian equivalence, budget balance, current account balance, short-run structural VAR.*

JEL CLASSIFICATION: F32; F41; H62; C32

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1. INTRODUCTION

The current account balance in Serbia underwent rapid change after the transformation process started following the regime switch in 2000. Serbia, with a huge delay compared to most of the CEE countries, started the transition process after spending the last decade of the last century in war conflicts and sanctions. This process is aimed at moving the country away from a state-regulated market to a market-oriented economy, putting the economy on the path of sustainable growth and development. On this path, in addition to a strong economic boom with rapid GDP growth and easy liquidity conditions, Serbia also experienced substantial trade deepening and large external imbalances. The current account deficit in 2004 reached 13.18% of GDP and increased to 19.87% in 2008 during the world financial crisis. The worsening of the current account deficit before the crisis was mainly caused by growth of investments in physical capital and booming domestic demand that was financed by large capital inflow due to expected future income. Hence, low private savings, accompanied by significantly large investments, worsened the trade balance. Even though after 2009 Serbia continued with rather lower current account deficits (about 5% of GDP on average), driven primarily by the slump in global and domestic demand in the post-crisis period, a new worsening of the current account is taking place, mainly due to the energy crisis caused by the war in Ukraine. This war has disrupted global energy markets, generating very high energy prices, and put all energy importers in an unfavorable position. Additionally, Serbia has failed to produce sufficient electricity for domestic demand, so electricity now has to be imported at very high prices.

Furthermore, Serbia may face significant increases in government spending in the short and medium term as a result of an increase in subsidies to cover the large disparity between high market prices for gas and electricity and the low prices paid by Serbian residents. This potential fiscal expansion could exacerbate the current account deficit even more. This has rekindled the debate whether budget deficits worsen the current account in Serbia or whether they have no impact. The existing literature distinguishes between two theoretical approaches. The first approach is associated with classical economists and refers to the theory of Ricardian equivalence (RE), according to which an increase in the budget deficit has no implications for the real economy (Barro, 1989). Namely, the decline in public savings, which is a consequence of tax cuts, is followed by an identical increase in private savings, which leaves national savings unchanged. In an open economy, there is no effect on the current account balance given that the rise in private savings is sufficient to avoid borrowing abroad. This approach assumes rational consumers who anticipate that the government will raise taxes in the future in order to be able to repay the accrued debt and meet its intertemporal budget constraint.

In contrast to RE, the conventional approach, which is based on models of overlapping generations with a finite time horizon, assumes that the growth of budget deficits (e.g. due to tax cuts) leads to an increase in both private savings and private consumption, thus reducing national savings (Diamond, 1965; Frenkel & Razin, 1992; Groth, 2015). In a small open economy, this leads to an increase in interest rates which, in the presence of relatively high capital mobility, results in a capital inflow into the country, then to exchange rate appreciation and finally to a deterioration of the current account. In a small open economy, with perfect capital mobility, even if interest rates do not rise, an increase in borrowing abroad (caused by smaller growth of private savings than the budget deficit) can worsen the current account. This approach is recognised in the literature as the twin deficit hypothesis (TDH) and is associated with the Keynesian group of economists.

Hence, the aim of this paper is to examine the impact of the budget balance on current account imbalances, apart from other determinants of the current account. The resulting current account deficits are likely to become a serious obstacle to maintaining external solvency in the case of a decrease in capital inflows (mainly FDI) due to a potential global economic recession. On top of this, additional fiscal expansion could significantly worsen the sustainability of the external position.

In previous relevant research, the TDH prevails, i.e., the budget deficit significantly affects the current account deficit. The literature is quite extensive and includes research that covers both developing and developed countries at the same time (Blanchard & Giavazzi, 2002; Mohammadi, 2004; Afonso et al., 2022). Examination of the TDH in Serbia has been almost non-existent in the literature so far. Tosun et al. (2014) considered Serbia within the framework of a study of a larger group of countries in the period from 2003 to 2010, but they did not find a long-run relation between the two deficits. On the other hand, Zildzović (2015) analysed the determinants of the current account in Serbia using model averaging techniques. He found that a reduction in the budget deficit has a positive effect on the improvement of the current account, which is in accordance with the TDH.

Our study aims to examine empirically dynamic relationships that exist between the budget deficit and the current account balance in Serbia using quarterly data in the period from 2005 to 2020. We used a multivariate vector autoregressive (VAR) model, along with the Granger causality test, decomposition of the error forecast variance and an impulse response function. In addition, in order to isolate current influences, the reduced form of the VAR model was transformed into a short-run structural vector autoregression (SVAR) model. The findings suggest that budget deficits lead to deterioration of the current account balance. This indicates that consumers in Serbia do not behave in a Ricardian way, which confirms the effectiveness of fiscal policy on the external balance. Nevertheless, the estimates of long-run Granger causality indicate evidence of mutual causality of the twin deficits in Serbia. In particular, a weak and statistically significant influence of the current account deficit of the previous period on budget deficits two periods later was also determined.

For the purpose of a robustness check, the generalised method of moments (GMM) was also applied due to high persistence of the current account. Thus, the current account lag value entered the set of explanatory variables, and we ended up with a dynamic model. However, since the Durbin–Wu–Hausman test showed that there is no endogeneity problem in the model, the estimates obtained by the OLS method can also be deemed relevant. These results also confirm the positive relation between these two balances and confirm the TDH in Serbia.

This study contributes to the existing literature in several aspects. First, a multivariate model for Serbia was employed instead of a bivariate model to avoid errors in establishing causality due to omitted relevant variables. Second, this study discovers the role of fiscal policy in managing Serbia's external balance, dispelling any doubt that fiscal expansion has nothing to do with current account balances.

The structure of this paper is as follows. Section 2 reviews the empirical literature. The model specification and descriptive data analysis are outlined in Section 3. Detailed econometric methodology is presented in Section 4, whilst the empirical results are reported in Section 5. Section 6 includes a robustness check, whereas the concluding observations are provided in Section 7.

2. LITERATURE REVIEW

In the existing literature, the TDH prevails, although the findings vary depending on the country, period of observation, and even the econometric techniques applied. The entire literature can be divided into two groups: the analysis of groups of countries and the analysis of individual countries. Empirical studies based on panel data analysis of a group of countries simultaneously model common and individual country characteristics and provide general results for the whole group of countries. On the other hand, studies on individual countries enable a deeper analysis using more flexible econometric research methods. Moreover, the research on the twin deficit hypothesis has undergone a gradual development of empirical methodology over time. In general, there is a tendency to use dynamic models due to the high persistence of both deficits, which requires the implementation of other econometric methods.

2.1. Analysis of a group of countries

Most previous research, when analysing groups of countries, has provided evidence of the TDH in both developed and developing countries (Afonso et al., 2022; Beetsma et al., 2008; Blanchard & Giavazzi, 2002; Mohammadi, 2004). Mohammadi (2004), using a sample of 20 industrial countries and 43 developing countries, showed that an increase in budget spending leads to a greater deterioration of the external deficit if it is financed by borrowing rather than by an immediate tax increase. In addition, Afonso et al. (2022) have shown that the TDH is valid, particularly with well-designed fiscal rules and the existence of an independent body (fiscal council), when the impact of the budget balance on the current account increases. One of the more comprehensive studies was conducted on a sample of 114 developing countries (1995–2015) by researchers from the IMF, Furceri & Zdzienicka (2020), who, using the VAR model in the panel, demonstrated that a sudden improvement in the budget balance of 1% of

GDP produces an improvement in the current account of 0.8% of GDP and thus confirmed the TDH.

The studies covering European countries seem to demonstrate certain inconsistencies. Several papers found no long-run relation between these two deficits, accepting the theory of RE and rejecting the TDH (Aristovnik & Djurić, 2010; Josifidis et al., 2021). On the other hand, Forte and Magazzino (2013) confirmed the TDH on a sample of 33 European countries (1970–2010), but only in countries facing high current account deficits (below -2% GDP). Two years later (Forte & Magazzino, 2015), these authors obtained the opposite results on a sample of EMU countries, using two different econometric techniques. Applying the Anderson-Hsiao IV estimators, they proved the TDH, whilst using the GMM approach, they confirmed the RE. Afonso & Coelho (2022) demonstrated that the positive impact of the budget balance on the current account balance is greater in countries outside the Eurozone, countries with large budget deficits, and countries with low exports. Their study was conducted on a sample of 28 EU countries in the period 1996–2019.

As for the panel data analysis of CEE countries, the literature is rather modest and the results vary widely. Ganchev et al. (2012) found a statistically significant but weak connection between these two deficits in CEE EU member countries in the period 1998–2009. The direction of influence is from the current account towards the budget balance. Reverse causality was also demonstrated by Obadic et al. (2014), who used Bulgaria, Romania, Poland and Croatia (1999–2011) as examples to show that in tax systems dominated by indirect taxes, the deterioration of the current account leads to growth in tax revenues due to increased imports. However, a study by Grubišić et al. (2018), which included all 16 CEE countries, did not obtain a statistically significant relation between these two balances.

2.2. Analysis of individual countries

Vector autoregression models (VAR) are commonly used along with the Granger causality test in the analysis of individual countries. Based on a detailed analysis of these papers, it is noted that all research in the literature relating to the analysis of individual countries was predominantly focused on determining the relation between these two deficits without including control variables, particularly in CEE countries.

Finally, an examination of twin deficits in Serbia is almost non-existent in the current literature. Tosun et al. (2014) included Serbia as part of their study and, based on quarterly data from 2003 to 2010, they did not find long-run causality between the two deficits. However, Zildzovic (2015), by analysing the determinants of the current account using model averaging techniques, found that the reduction of the budget deficit has a positive effect on the improvement of the current account, which is consistent with the TDH.

Given the large number of papers analysing individual countries, a detailed review of the literature is provided in Table 1.

The empirical literature reviewed above focuses primarily on developed countries and the TDH dominates over RE theory. Developing countries are mostly considered to be part of a larger group of countries and the results are not clear-cut. Countries with higher current account deficits tend to support the TDH more strongly, while over-indebted countries confirm the RE theory. Furthermore, the group analysis of European countries yielded different outcomes depending on the econometric techniques used, the current account balances, and Eurozone membership. The most recent findings are based on a dynamic panel data analysis of the current account balance and the budget balance, including other current account determinants. Considering individual countries, which are the subject of this study as well, the majority of empirical studies employed cointegration analysis and Granger causality tests. Most of the studies specified a bivariate VAR model, omitting important determinants of the current account balance. As with the group of countries, the prevailing results are that the budget deficit significantly affects the current account deficit, refuting the Ricardian equivalence theory. Finally, the studies covering CEE countries produce a variety of results, primarily depending on the time period examined. In several CEE countries where indirect taxes dominate, a negative relationship between the two balances was even observed, as an increase in exports resulted in a tax rise.

Table 1 Literature review of TDH (analysis of individual countries)

Reference	Sample	Control variables	Econometric technique	Conclusion
All countries (mixed)				
Khalid and Guan (1999)	Developed countries (US, UK, France, Canada and Australia) Developing countries (India, Indonesia, Pakistan, Egypt and Mexico) 1955–1993	None	Cointegration analysis	TDH (US, France, Egypt, Mexico) RE (UK, Austrlia) RC (Indonesia, Pakistan) Bi-directional (Canada, India)
Piersanti (2000)	OECD countries 1970–2007	None	Granger causality tests and GMM	TDH
Kouassi et al. (2004)	20 developed and developing countries 1969–1998	None	Granger causality (Toda&Yamamoto approach)	TDH (Italy, Israel) RE (others)
Salvatore (2006)	G-7 countries 1973–2005	Real GDP growth rate	Dynamic regression	TDH
S. Kim & Roubini (2008)	US 1973–2004 quarterly data	Real GDP, 3-month real interest rate, real exchange rate	VAR model, impulse response function	TDH divergence
Daly & Siddiki (2009)	23 OECD countries	Real interest rate	Cointegration analysis with regime shift	TDH in 13 countries Without regime switch (significantly less)
Asian countries				
Vamvoukas (1999)	Greece	None	Cointegration analysis and	
C.-H. Kim & Kim (2006)	Korea 1970–2003	None	Granger causality (Toda&Yamamoto approach)	Reverse causality (from current account to fiscal balance)
Baharumshah et al. (2006)	Malaysia, Indonesia, Philippines, Thailand 1976–2000 quarterly data	None	VAR model, Variance decomposition, Granger causality	TDH Bi-directional (Malaysia, Philippines) Reverse causality (Thailand)
Ogbonna (2013)	Nigeria 1960–2011	Real lending interest rate, real GDP growth, openess, real exchange rate	Multivariate VAR	TDH in long run No relation in short run
European countries				
Vamvoukas (1999)	Greece	None	Cointegration and ECM	TDH
Magazzino (2012)	Italy 1970–2010	None	Granger causality	Reverse causality only in short run
Algieri (2013)	GIIPS 1980–2012 quarterly data	None	Granger causality (Toda&Yamamoto approach)	RE
Trachanas & Katrakilidis (2013)	GIIPS 1971–2009	None	Regime switch model	TDH
Nikiforos et al. (2015)	Greece 1980–2010 quarterly data	None	Granger causality	TDH both in short and long run Reverse causality
CEE countries				
Obadic et al. (2014)	Bulgaria, Croatia, Poland, Romania 1999–2011 quarterly data	None	VAR	TDH divergence due to indirect taxes
Tosun et al. (2014)	Bulgaria, Latvia, Lithuania, Poland, Romania, Serbia, Slovenia 1990–2013	None	ARDL model	TDH only in Bulgaria
Turan & Karakas (2018)	Czech Republic, Hungary, Slovakia, Poland, Romania, Slovenia, Croatia 1999–2016 quarterly data	None	NARDL model	TDH in Czech Republic, Hungary, Slovakia Reverse causality in Poland and Romania TDH only in short run in Croatia
Matę (2019)	Poland, Czech Republic, Hungary	None	VAR and cointegration test	TDH in Czech Reverse causality in Hungary

3. MODEL SPECIFICATION AND DATA

This study tests the relation between the current account balance and the government budget balance. The starting point is a well-known accounting identity of the gross national product, as shown in Equation (1). On the left-hand side we have gross national product (GNP) as the sum of income derived from producing goods and services for private consumption *(C)*, private investment *(I)*, public goods and services *(G)*, and exports *(X)*. Imports *(Z)* are included as negative item to avoid double accounting. ($X - Z$) refers to the trade balance plus net factor income. On the right-hand side, we have possible uses of GNP: it can be consumed *(C)*, saved *(S)*, paid as taxes *(T)*, or transferred abroad *(Tr)*.

$$C + I + G + (X - Z) = C + S + T + Tr \tag{1}$$

Rearranging this equation results in Equation (2), in which the expression on the left-hand side refers to the current account balance ($X - Z - Tr$), while $T - G$ refers to the budget balance (*BB*) or government savings (S^G), *S* denotes private savings and *I* private investments. Overall, current account balance *(CA)* is equal to national savings (government plus private) minus private investments (Equation (3))

$$(X - Z) - Tr = (T - G) + (S - I) \tag{2}$$

$$CA = BB + S - I \tag{3}$$

According to RE, a decrease in government savings (*BB*) due to tax cuts leads to an increase of equal magnitude in private savings (*S*), leaving *CA* unchanged. In contrast to RE, the TDH asserts that a decrease in government savings primarily increases private consumption, but partly increases private savings as well. As a result, private savings increase by an amount that is smaller than that of the initial tax cut, and national savings decline, which worsens the current account.

An empirical model testing the twin deficit hypothesis is presented in Equation (4), where CA_t is the current account at time t (t = 1,...,T), BB_t is the consolidated budget balance or government savings ($T - G$), Z_t represents a set of control variables (other determinants of the current account), α is a constant, whilst ε_t is the error term.

$$CA_t = \alpha + \beta BB_t + \theta Z_t + \varepsilon_t \tag{4}$$

The twin deficits hypothesis predicts that a worsening of the budget balance leads to a worsening of the current account balance ($\beta > 0$), while the Ricardian equivalence theory predicts that there is no impact on the current account balance ($\beta = 0$).

As the current account balance shows high persistence over time, this implies including its lagged value in the set of explanatory variables. Therefore, the static model becomes the dynamic model and is shown in Equation (5).

$$CA_t = \alpha + \gamma CA_{t-1} + \beta BB_t + \theta Z_t + \varepsilon_t \tag{5}$$

This study is based on quarterly data in the period 2005–2020. Before 2005, time series of Serbian data either do not exist or are not reliable. The dependent variable in the regression analysis is the ratio of the current account balance to GDP. Fiscal policy is covered through the ratio of the consolidated budget balance to GDP. Most of the data are taken from the National Bank of Serbia (NBS), the Statistical Office of the Republic of Serbia (SORS), and the Ministry of Finance (MF). As expected, most time series have a pronounced seasonal pattern, so the data are initially adjusted for the seasonality by applying the TRAMO-SEATS procedure (Gómez & Maravall, 1996). All the variables are presented in Table 2.

Table 2 List of all the variables used in the analysis

Variables	Full name and description	Expressed season	Source
CA	Current account (% of GDP)	YES	NBS
BB	Consolidated budget balance (% of GDP)	YES	MF
RER	Real effective exchange rate, change	NO	NBS
TT	Terms of trade – index of net export prices (ratio between export and import prices, individual products weighted by the share of net exports in GDP)	NO	SORS
GNI	Gross national income p.c. relative to EMU	YES	SORS and Eurostat
INVP	Private investments (% of GDP)	YES	SORS and MF

In addition to the basic variables, other determinants of the current account are also included. First of all, we include the real exchange rate. The impact of the change in the real exchange rate on the current account is ambiguous. Immediately after real depreciation of the currency's value, exports immediately become cheaper and imports more expensive, leading to a deterioration in the trade balance. Shortly thereafter, the volumes of exports start to increase steadily due to cheaper prices while the volumes of imports decrease, leading to an improvement of the trade balance. In the literature, this is well known as the J-curve effect.

The terms of trade are included in the model to capture the effects of export and import price movements on the current account. The effects of the stage of economic development are measured by the gross national income (GNI) per capita relative to the average GNI per capita of the 27 EMU members. Furthermore, the model includes private investments, which are expected to have a negative impact on the current account.

As a preliminary analysis, descriptive statistics are presented in Table 3. On average, the quarterly current account deficit in Serbia was 8.39%, with it reaching a maximum of 24% during the global financial crisis, whilst the lowest deficit level was 2.15% in the second quarter of 2015. After 2000, Serbia's current account deficit tended to deteriorate, then stabilising after the crisis (Figure 1). Namely, after 2000, when the transition process began in Serbia, a large inflow of foreign capital took place in the form of loans and foreign direct investments, which also affected the development and efficiency of the financial market. Nonetheless, this huge inflow of capital also caused a huge growth in domestic demand (both consumption and investment) and with extremely low savings rates the current account balance further deteriorated. Moreover, the Serbian economy was drained by spending on the war conflicts and sanctions of last decade of last century. After the global financial crisis in 2008, capital inflows and domestic demand dropped significantly, while savings rates started to rise gradually, which led to the reduction of the current account deficit.

The consolidated budget balance recorded an average quarterly deficit of 2.50% of GDP in the period from 2005 to 2020. The largest deficit, of 9.42%, was reached in the last quarter of 2014, after which the process of fiscal consolidation began and Serbia recorded a surplus of 3.10% in the third quarter of 2017.

Table 3 Descriptive statistics of all variables

Variables	Average	Median	Stand. deviation	Minimum	Maximum
CA	**-8.39**	**-6.99**	**5.05**	**-23.99**	**-2.15**
BB	**-2.50**	**-2.96**	**3.32**	**-9.42**	**3.10**
RER	0.40	-0.03	3.19	-8.72	8.03
TT	102.41	102.3	5.38	89.20	114.8
GNI	0.36	0.37	0.03	0.27	0.43
INVP	15.97	15.44	2.25	12.31	21.39

Figure 1 Current account and budget balance in Serbia (quarterly data)

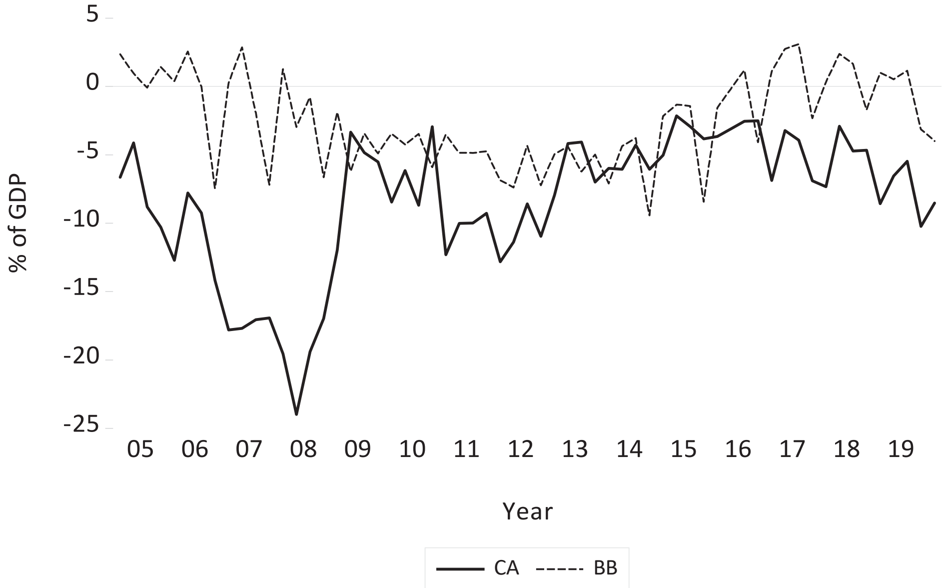


Figure 1 clearly shows the trend separation between these two deficits before the crisis and particularly during the crisis, whereas in the post-crisis period they return to the same trend line. Hence, the overall correlation rate is only 0.04. As

Tosun et al. (2014) only covered the pre-crisis and crisis period (2003–2010), they failed to identify a long-run linkage between these two deficits in Serbia.

4. ECONOMETRIC METHODOLOGY

In the existing literature, the analysis of the TDH in individual countries is mostly based on bivariate vector autoregression (VAR) models with the implementation of cointegration analysis and Granger causality tests (Kim & Roubini, 2008; Matę, 2019; Obadic et al., 2014; Ogbonna, 2013). Furthermore, only one study examines the TDH in Serbia (Tosun et al., 2014). However, this study employed a bound testing approach used an autoregressive distributed lag (ARDL) model and analysed quarterly data for 7 CEE countries, including Serbia. They failed to identify any long-run relationship between the two deficits, but the results might be biased due to the omission of relevant variables. Therefore, in testing the causality between the budget and current account deficits, our model also includes additional *CA* determinants. Moreover, given that the real effective exchange rate (*RER*) is potentially endogenous, this requires the application of a multivariate vector autoregression model, which is shown in Equation (6).

$$CA_t = a_1 + \varphi_{1i}^1 CA_{t-i} + \varphi_{1i}^2 BB_{t-i} + \varphi_{1i}^3 RER_{t-i} + e_1 Z_t + d_1 V + u_{1t} \tag{6}$$

$$BB_t = a_2 + \varphi_{2i}^1 CA_{t-i} + \varphi_{2i}^2 BB_{t-i} + \varphi_{2i}^3 RER_{t-i} + e_2 Z_t + d_2 V + u_{2t}$$

$$RER_t = a_3 + \varphi_{3i}^1 CA_{t-i} + \varphi_{3i}^2 BB_{t-i} + \varphi_{3i}^3 RER_{t-i} + e_3 Z_t + d_3 V + u_{3t}$$

Vector Z_t includes other control variables that are exogenous: terms of trade (*TT*), private investments (*INVP*), and relative GNI (*GNI*), whilst vector *V* includes dummy variables that aim to neutralise the impact of structural breaks.

Based on the multivariate VAR model, various econometric techniques were applied, such as the Granger causality test. An impulse response function was used to produce the time path of the current account balance when an unexpected shock to the budget balance occurred. Along with the impulse response function, the forecast error variance decomposition was also calculated to see the contribution to the forecast error variance from a specific exogenous shock.

It is well-known that equations in a reduced VAR model contain the variable's lagged values and lagged values of other variables (Equation (6)) and do not give any information about the impact of a direct change in one variable on other variables in the model. In order to obtain the impact of the level values of the budget balance on the current account balance, the standard VAR model was transformed into a short-run structural vector autoregression (SVAR) model. A short-run SVAR model involves a recursive setup identified by short-run restrictions on the impact effects of the structural shocks. Mostly, it is used when the model is partially identified with only one structural shock of interest, as in this study. After estimating the reduced VAR model to calculate the MM estimator of structural parameters, a Cholesky decomposition was applied. Finally, the recursive form of the short-run structural VAR model, as shown in Equation (7), was obtained.

$$RER_t = a_{11}^{i} RER_{t-i} + a_{12}^{i} BB_{t-i} + a_{13}^{i} CA_{t-i} + \varepsilon_{1t} \tag{7}$$

$$BB_t = a_{21}^{0} RER_t + a_{21}^{i} RER_{t-i} + a_{22}^{i} BB_{t-i} + a_{23}^{i} CA_{t-i} + \varepsilon_{2t}$$

$$CA_t = a_{31}^{0} RER_t + a_{32}^{0} BB_t + a_{31}^{i} RER_{t-i} + a_{32}^{i} BB_{t-i} + a_{33}^{i} CA_{t-i} + \varepsilon_{3t}$$

The third equation is the focus of this paper: the a_{32}^{0} coefficient indicates the impact of the budget balance on the current account balance. In countries where the TDH holds, the a_{32}^{0} coefficient is expected to be statistically greater than 0. However, RE theory predicts the a_{32}^{0} coefficient is close to 0.

5. EMPIRICAL RESULTS

In order to apply any of the foregoing methods, we must determine the level of integration of each of the time series. For this purpose, standard unit root tests were applied: the augmented Dickey–Fuller (ADF) test and the Kwiatkowski–Phillips–Schmidt–Shin (KPSS) test. It should be noted that Serbia, like other countries worldwide, was affected by the financial crisis in 2008, causing significant structural changes in its economy. It is commonly known that the ADF test gives biased estimates against rejecting the null hypothesis of the existence of a unit root. Therefore, the Zivot-Andrews (ZA) unit root test with a structural break was also applied. This test endogenously estimates break points.

The results are shown in Table 4 and indicate that all series are stationary. As for the current account, due to the presence of a structural break, the ADF test demonstrated that the series has a unit root, but both the KPSS test and the ZA test indicated strict stationarity of this series. In addition, the budget balance is a trend stationary series, which has a structural break in the intercept. Therefore, when the standard ADF and KPSS test were applied, the result is that it has a unit root. However, on applying the ZA test, it was found that the series is highly stationary. Moreover, when we conducted standard ADF testing on two subperiods, stationarity is also evident. The situation is similar with private investments and relative GNI.

Table 4 Results of the unit root test

Variables	In levels		
	ADF[1]	KPSS[2]	ZA test[1]
CA	**-3.13(0)**	**0.08**	**-4.25***(0)**
BB	**-2.05(1)**	**0.22*****	**-3.53***(2)**
INVP	-2.01(3)	0.24	-3.77**(2)
GNI	-1.98(0)	0.40*	-3.40**(3)
RER	-6.21***(1)	0.18	/
TT	-4.52***(0)	0.20	/

[1]H_0: The series has a unit root, [2]H_0: The series is stationary

*, **, and *** indicate a p-value at the 10%, 5% and 1% levels, respectively

In the next step, a multivariate VAR model with three equations was estimated. To select the number of lags in the VAR model specification, Akaike and Schwartz information criteria (AIC and SC, respectively) were first used. Given that these criteria suggested a different number of lags in the model (AIC proposes the specification of a VAR model that includes 3 lags, whereas SC suggests one lag), Sim's modified likelihood ratio test (LRT) was used, confirming the result of SC. Finally, the VAR (1) model specification of dimension 3 was estimated. As already noted in the methodology, in addition to other exogenous variables, three dummy variables were included (V_1 and V_2 include the period of the financial crisis, whilst V_3 includes the onset of fiscal consolidation). V_1 takes the value 1 in the second quarter of 2008, whereas it takes 0 otherwise, V_2 takes the value 1 in the fourth quarter of 2008, and 0 otherwise, and V_3 takes the value 1 in the fourth quarter of 2012 and first quarter of 2013, and 0 for the rest.

The results of the specification tests are shown in Table 5. A p-value greater than 5% for the first four lags (given that they are quarterly data) indicates that there is no serial autocorrelation of the residuals. To examine the distribution of the residuals, the Doornik–Hansen normality test was applied, according to which the residuals are fairly normally distributed. Furthermore, the values of all 3 roots of characteristic polynomial are strictly less than 1, so that the multivariate VAR (1) model fulfills the stability condition.

Table 5 Tests of VAR (1) model specification

Variables	**Coefficient**	**P - value**
Autocorrelation test (LM test)		
AR(1)	6.7120	0.667
AR(2)	5.2355	0.813
AR(3)	15.0108	0.091
AR(4)	6.4397	0.695
Normality test (Doornik–Hansen test)	5.0533	0.537
Stability condition (values of characteristic roots)	**0.6425 0.6425 0.0408**	

The Granger causality test showed bi-directional causality between the current account and the budget balance (Table 5). It is important to point out that this is a statistical concept of causality based on prediction, i.e., the influence of one variable in period *t-1* on another variable in period *t*. Besides significant causality from the current account to the budget balance, the Granger causality test also shows significant causality from the budget balance to the current account balance.

Table 6 Granger causality test

Variables	**χ^2 stat**	**p - value**
Current account (CA)		
Budget balance (BB) does not cause CA	4.5314**	0.033
The real exchange rate (RER) does not cause CA	10.2915***	0.001
Budget balance (BB)		
Current Account (CA) does not cause BB	6.6638**	0.010
Real exchange rate (RER) does not cause BB	0.9263	0.335
Real exchange rate (RER)		
Current account (CA) does not cause RER	1.6830	0.194
Budget balance (BB) does not cause RER	5.6349**	0.018

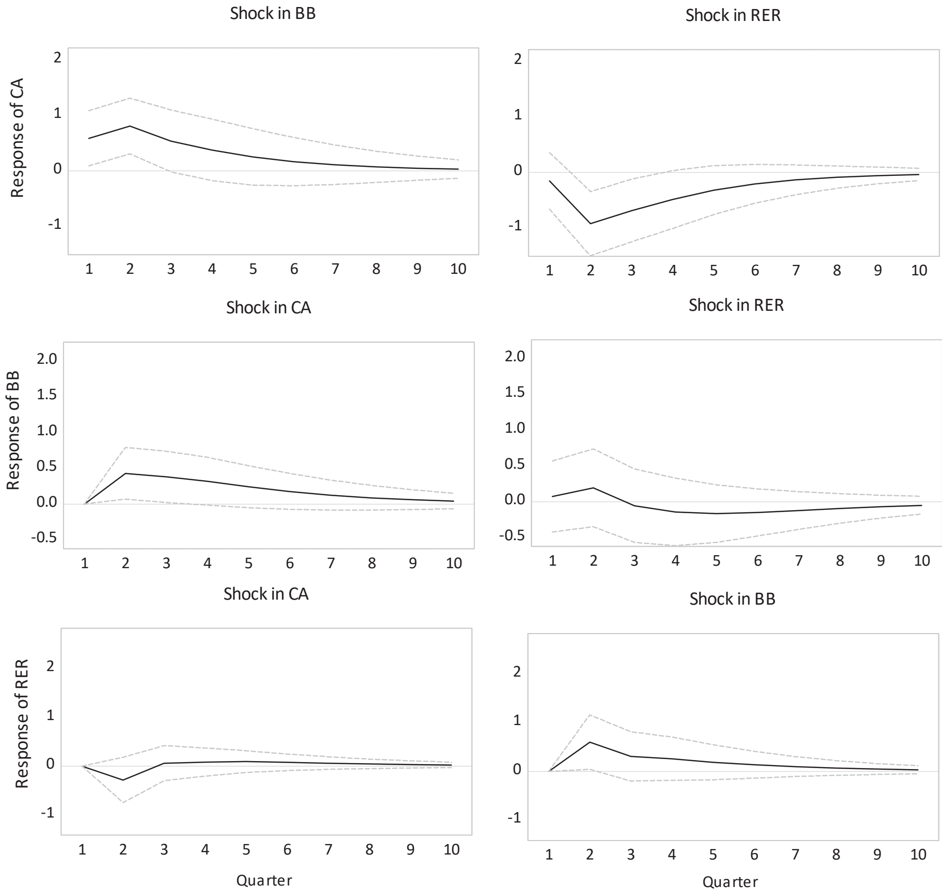
***$p < 0.01$, ** $p < 0.05$, * $p < 0.10$

It is relevant to determine the sign of the reaction, which is achieved by applying the impulse response function (Figure 2). It is clearly seen that an unexpected increase in the budget deficit creates a statistically significant positive response of the current account deficit, which is in line with the TDH. The response is significant and occurs immediately after the first quarter; then, in the second quarter, the response is even greater, and after that it persists for 2 years.

The budget balance in Serbia responds in a similar way to the current account shock but with lower intensity. Due to an unexpected increase in the current account deficit, the budget balance deficit increases, but only after two quarters, and then gradually decreases.

A shock to the real exchange rate registers the largest response of the current account. The strongest negative response is after two quarters and remains significantly negative until the end of the fourth quarter. After the first quarter, the response of the current account is also negative, but almost negligible. This delayed negative impact of the real exchange rate on the current account is the J-curve effect mentioned earlier: the current account balance initially worsens following the currency depreciation, and then quickly recovers and finally surpasses its previous decline. It is interesting to note that the budget balance shock has a statistically significant effect on the real exchange rate, but only in the second quarter.

Figure 2 Estimation of ordinary impulse response function based on Cholesky one-standard-deviation shock



Moreover, these results are supported by decomposition of the error forecast variance presented in Table 7. It shows that an exogenous shock of the budget balance accounts for 15% in forecasting the current account variations after three quarters, whereas that of the real exchange rate accounts for 17% after four quarters.

Table 7 Variance decomposition of forecast errors

Quarter	CA	BB	RER	SUM
1	89.49	7.67	2.83	100
2	71.47	13.78	14.74	100
3	67.40	14.62	17.98	100
4	65.93	14.73	19.33	100
5	65.36	14.70	19.93	100
6	65.13	14.66	20.20	100
7	65.04	14.64	20.31	100
8	65.01	14.63	20.36	100
9	64.99	14.62	20.38	100
10	64.99	14.62	20.39	100

To obtain the direct effects of the budget balance on the current account, the reduced form of the multivariate VAR (1) model was transformed into a short-run SVAR model (Equation (7)) and the results are shown in Table 8. The table shows that the coefficient a_{32}^{0} is positive and statistically significant at the 5% level, which indicates that the budget balance significantly affects the current account balance. This result confirms the TDH and rejects RE theory in Serbia. More precisely, an increase in the budget deficit (as % of GDP) of 1 percentage point leads to a worsening of the current account balance (as % of GDP) of 0.31 percentage point. The result indicates that current budget deficits are accrued at the public debt account and are transferred to the future generations, which is contrary to the RE theory.

The impact of the real exchange rate on the current account has a negative sign and amounts to -0.14, but it is statistically insignificant (p-value = 0.18). As already pointed out in the text, the influence of the real exchange rate has a delayed effect on the current account (J-curve effect). Namely, the temporary exchange rate appreciation immediately tends to improve the current account deficit as it improves the terms of trade. But when export volumes start to decrease, and import volumes increase, the current account worsens and even surpasses the initial improvement. This is in line with J-curve effect. Due to these two opposing effects, initially the impact of the exchange rate may be positive or statistically insignificant, but later a statistically significant and negative impact on the external balance is expected.

Table 8 Short-run structural VAR model for Serbia

	RER	BB	CA
RER	1	0	0
BB	0.0114	1	0
CA	-0.1422	0.3065**	1

***p < 0.01, ** p < 0.05, * p < 0.10

6. ROBUSTNESS CHECK

For the robustness check, we applied a two-stage least squares (2SLS) estimator, which is a special case of the general method of moments (GMM), due to the current account persistence and potential endogeneity problem (Table 9). Only the first lag of the current account is statistically significant at the 5% level, whilst the second is not. Therefore, persistence can be observed, but not to a great extent. Besides the lagged value of the current account, a change in the real exchange rate appears as a potentially endogenous variable. The external deficits directly influence exchange rate variations. So, the real exchange rate entered the model with one lag in order to avoid potential endogeneity, and also due to the delayed expected effect on the current account.

Table 9 Twin deficit model for Serbia – 2SLS method

Variables	Coefficient	Standard error	P - value
CA_{t-1}	0.3675***	0.1262	0.005
BB_t	**0.3774*** **	**0.1307**	**0.008**
RER_{t-1}	-0.3457***	0.0837	0.000
$INVP_t$	-1.2098***	0.2572	0.000
GNI_t	21.4352***	6.9626	0.003
TT_t	0.0719**	0.0352	0.046
V_1[1]	-8.0612***	2.1770	0.000
V_2[2]	-5.6857***	1.7972	0.003
V_3[3]	4.8371**	1.8793	0.013
Specification tests			
R^2	0.9006		
Endogeneity test (Durbin-Wu-Hausman test)	**0.249**		**0.618**
Hansen's J statistic on overidentification	0.043		0.836
Cragg-Donald statistics, Stock&Yogo crit. value	17.41		

***p < 0.01, ** p < 0.05, * p < 0.10

[1]V_1 is dummy variable that takes the value of 1 for Q2 2008, and 0 otherwise. This is the period of the financial crisis.

[2]V_2 is dummy variable that takes the value of 1 for Q1 2011, and 0 otherwise. It is a period of correction of the current deficit primarily as a result of the slowdown in investments, and thus the reduction in imports of intermediate products.

[2]V_3 is a dummy variable that takes the value of 1 for Q1 2013, and 0 otherwise. This is the commencement of the period of fiscal consolidation.

However, Durbin-Wu-Hausman test statistics clearly demonstrate that no problem pertaining to the endogeneity of the model is observed (Table 9). Hence, the null hypothesis that the OLS method gives consistent and efficient estimations cannot be rejected. Therefore, estimates based on the OLS method are also relevant (Table 10) and the coefficient estimates that arise by using the two methods are similar. The Ricardian equivalence theory is rejected at the 1% significance level as β is 0.34, which implies that if the consolidated budget deficit as a percentage of GDP increases by 1 percentage point, the current account deficit as a percentage of GDP will rise by 0.34 percentage points. The result is fully in line with the short-run SVAR approach (the coefficient with a budget balance was 0.31) and confirms the TDH.

Table 10 Twin deficit model for Serbia - OLS method

Variables	**Coefficient**	**Standard error**	**P - value**
CA_{t-1}	0.4155***	0.0811	0.000
$\mathbf{BB_t}$	**0.3421*** **	**0.1095**	**0.003**
RER_{t-1}	-0.3321***	0.0788	0.000
$INVP_t$	-1.1217***	0.1857	0.000
GNI_t	21.1002***	6.9058	0.004
TT_t	0.0624**	0.0294	0.039
V_1[1]	-8.0793***	1.8296	0.000
V_2[2]	-5.8040***	1.7751	0.002
V_3[3]	4.9151**	1.8662	0.036
Specification tests			
R^2	0.9013		
Normality test (Jarque-Bera test)	1.294		0.524
Autocorrelation test (Breusch-Godfrey test)	0.626		0.646
Heteroscedasticity test (Breusch-Pagan-Godfrey test)	0.468		0.889

***p < 0.01, ** p < 0.05, * p < 0.10

[1]V_1 is dummy variable that takes the value of 1 for Q2 2008, and 0 otherwise. This is the period of the financial crisis.

[2]V_2 is dummy variable that takes the value of 1 for Q1 2011, and 0 otherwise. It is a period of correction of the current deficit primarily as a result of the slowdown in investments, and thus the reduction in imports of intermediate products.

[2]V_3 is a dummy variable that takes the value of 1 for Q1 2013, and 0 otherwise. This is the commencement of the period of fiscal consolidation.

Other variables also have a significant effect on the current account. Firstly, the appreciation of the real exchange rate one period back leads to a deterioration of the current account deficit (-0.33). Furthermore, the capital production coefficient in the private sector is highly statistically significant and negative, which implies that the increase in private investment leads to a rise in the current account deficit. This is the expected result given that it is a small open economy in which private savings are quite low. A positive coefficient of the relative income per capita variable indicates that if relative income is below the average, it will be associated with a current account deficit, and if relative income is above the average, it will be associated with a surplus. More impecunious countries usually have lower savings than investments and a relatively high rate of returns to investments. Thus, a country anticipating large future income increases consumption by borrowing now and repaying later.

Furthermore, the regression results indicate that the improvement in the terms of trade has a positive effect on the current account. This is in line with the Harberger–Laursen–Metzler effect, which predicts a positive relation between temporary changes in the terms of trade and national savings through consumption smoothing. Namely, due to the deterioration of the terms of trade, there is a decrease in the current real income, which is greater than the decrease in the permanent income of individuals. Given that the marginal propensity to consume is less than 1, a drop in national consumption is predicted, but also a decline in national savings, which affects the increase in the external account deficit. Alternatively, the deterioration of the terms of trade directly worsens the current account balance, and its improvement enhances the current account balance. The output gap and the country's openness to trade have the expected signs, but they are not statistically significant in the model for Serbia and are excluded from the model.

It is important to note that all the tests of the OLS model specification are satisfied. First of all, the Jarque–Bera test confirms that the residuals are normally distributed. The Breusch–Godfrey autocorrelation test confirms that there is no serial correlation between the residuals, whereas according to the Breusch–Pagan–Godfrey test statistic, one cannot reject the null hypothesis of homoscedasticity.

Regarding the GMM/IV approach, in addition to the endogeneity test, other diagnostic tests were also conducted. Hansen's test of overidentification is satisfied (the resulting J statistics fail to reject its null), as are the Cragg–Donald statistics, which reject the weak instrument hypothesis with a maximum bias (at 15% significance).

7. CONCLUSION

This paper examines the empirical linkage between the budget deficit and current account imbalances in Serbia based on quarterly data in the period from 2005 to 2020. For this purpose, the following econometric techniques were used: the vector autoregression model (VAR) and the structural vector autoregression model (SVAR) in order to extract current influences. To check the robustness, the OLS method was applied, as well as the generalised method of moments (GMM) due to the high persistence of the current account. The results of all the models unequivocally suggest that the budget deficit affects the current account deficit, rejecting the hypothesis that consumers in Serbia behave Ricardian and confirming the TDH. It is interesting that the Granger causality test showed that this relation is bi-directional, i.e., in addition to the fact that the previous values of the budget balance positively affect the current account, the previous values of the current account also significantly and positively affect the budget balance but to a much lower extent.

This paper contributes to the existing literature in two ways. Firstly, this research provides a detailed and comprehensive analysis and with various econometric techniques confirms the twin deficit hypothesis in Serbia. In the previous literature, Serbia was included in the analysis of a larger group of countries and no statistically significant connection was found between these two balances (Tosun et al., 2014). Moreover, other control variables were not included in that analysis.

Secondly, this research has a practical application. Such research findings remove any doubts about the effectiveness of using fiscal policy to manage the external balance and suggest that this fiscal policy should be aimed at mitigating the current account deficit in Serbia. The findings clearly demonstrate that an increase in budget deficits significantly worsens the current account. This finding implies that the Serbian government should end the policy of extraordinarily large deficits started by the COVID crisis in 2020. Instead, Serbia should engage in tightening fiscal policy in the following period to curb further deterioration of the current account deficit, which is caused primarily by the energy crisis due to the war in Ukraine. Additional fiscal expansion could significantly worsen the sustainability of the external position. In the long run, macroeconomic policymakers should resort to policies that encourage fiscal consolidation to rectify external account imbalances.

REFERENCES

Afonso, A., & Coelho, J. C. (2022). The role of fiscal policies for external imbalances: Evidence from the European Union. *Economic Systems*, *46*(2), 100985. https://doi.org/10.1016/j.ecosys.2022.100985

Afonso, A., Huart, F., Tovar Jalles, J., & Stanek, P. (2022). Twin deficits revisited: A role for fiscal institutions? *Journal of International Money and Finance*, *121*, 102506. https://doi.org/10.1016/j.jimonfin.2021.102506

Algieri, B. (2013). An empirical analysis of the nexus between external balance and government budget balance: The case of the GIIPS countries. *Economic Systems*, *37*(2), 233–253. https://doi.org/10.1016/j.ecosys.2012.11.002

Aristovnik, A., & Djurić, S. (2010). Twin Deficits and the Feldstein Horioka Puzzle: A Comparison of the EU Member States and Candidate Countries. *Munich Personal RePEc Archive*, *Paper No. 24149*, 21. https://mpra.ub.uni-muenchen.de/24149/

Baharumshah, A. Z., Lau, E., & Khalid, A. M. (2006). Testing Twin Deficits Hypothesis using VARs and Variance Decomposition. *Journal of the Asia Pacific Economy*, *11*(3), 331–354. https://doi.org/10.1080/13547860600764245

Barro, R. J. (1989). The Ricardian Approach to Budget Deficits. *Journal of Economic Perspectives*, *3*(2), 37–54. https://doi.org/10.1257/jep.3.2.37

Beetsma, R., Giuliodori, M., & Klaassen, F. (2008). The Effects of Public Spending Shocks on Trade Balances and Budget Deficits in the European Union. *Journal of the European Economic Association*, *6*(2–3), 414–423. https://doi.org/10.1162/JEEA.2008.6.2-3.414

Blanchard, O., & Giavazzi, F. (2002). Current Account Deficits in the Euro Area: The End of the Feldstein-Horioka Puzzle? *Brookings Papers on Economic Activity*, *2002*(2), 147–186.

Daly, V., & Siddiki, J. U. (2009). The twin deficits in OECD countries: Cointegration analysis with regime shifts. *Applied Economics Letters*, *16*(11), 1155–1164. https://doi.org/10.1080/13504850701349179

Diamond, P. A. (1965). National Debt in a Neoclassical Growth Model. *The American Economic Review*, *55*(5), 1126–1150.

Forte, F., & Magazzino, C. (2013). Twin Deficits in the European Countries. *International Advances in Economic Research*, *19*(3), 289–310. https://doi.org/10.1007/s11294-013-9406-3

Forte, F., & Magazzino, C. (2015). Ricardian equivalence and twin deficits hypotheses in the euro area. *Journal of Social and Economic Development*, *17*(2), 148–166. https://doi.org/10.1007/s40847-015-0013-4

Frenkel, Jacob A. & Razin, Assaf. (1992). Fiscal Policies and the World Economy. Cambridge, MA: MIT Press.

Furceri, D., & Zdzienicka, A. (2020). Twin Deficits in Developing Economies. *Open Economies Review*, *31*(1), 1–23. https://doi.org/10.1007/s11079-019-09575-1

Ganchev, G. T., Stavrova, E., & Tsenkov, V. (2012). Testing the Twin Deficit Hypothesis: The Case of Central and Eastern Europe Countries. *International Journal of Contemporary Economics and Administrative Sciences*, *2*(1), 1–21

Gómez, V., & Maravall, A. (1996). Programs TRAMO and SEATS, Instruction for User (Beta Version: September 1996). (*Working Paper*). *Banco de España*. https://econpapers.repec.org/paper/bdewpaper/9628.htm

Groth, C. (2015). Lecture Notes in Macroeconomics. (Chapter 6). https://web2.econ.ku.dk/okocg/MAT-OEK/Mak%C3%98k2/Mak%C3%98k2-2015/Forel%C3%A6sninger/Pensum-M2-i-een-fil-2015.pdf

Grubišić, Z., Kamenković, S., & Zdravković, A. (2018). Impact of government balance and exchange rate regime on current account during the economic cycle: Evidence from CEE countries. *Zbornik Radova Ekonomskog Fakulteta u Rijeci: Časopis Za Ekonomsku Teoriju i Praksu*, *36*(1), 307–334. https://doi.org/10.18045/zbefri.2018.1.307

Josifidis, K., Dragutinović-Mitrović, R., & Bodor, S. (2021). The effect of fiscal deficits on the external imbalances in the European Union. *Panoeconomicus*, *68*(5), 625–652. https://doi.org/10.2298/PAN190523017J

Khalid, A. M., & Guan, T. W. (1999). Causality tests of budget and current account deficits: Cross-country comparisons. *Empirical Economics*, *24*(3), 389–402. https://doi.org/10.1007/s001810050062

Kim, C.-H., & Kim, D. (2006). Does Korea have twin deficits? *Applied Economics Letters*, *13*(10), 675–680. https://doi.org/10.1080/13504850500404910

Kim, S., & Roubini, N. (2008). Twin deficit or twin divergence? Fiscal policy, current account, and real exchange rate in the U.S. *Journal of International Economics*, *74*(2), 362–383. https://doi.org/10.1016/j.jinteco.2007.05.012

Kouassi, E., Mougoué, M., & Kymn, K. O. (2004). Causality tests of the relationship between the twin deficits. *Empirical Economics*, *29*(3), 503–525. https://doi.org/10.1007/s00181-003-0181-5

Magazzino, C. (2012). The Twin Deficits Phenomenon: Evidence from Italy. (*SSRN Scholarly, Paper No. 2333444*). https://papers.ssrn.com/abstract=2333444

Matę, A. (2019). Interactions between the budget and the current account balance: Twin deficits in selected Central and East European countries. *Journal of Management and Financial Sciences*, *38*, 47–60. https://doi.org/10.33119/JMFS.2019.38.4

Mohammadi, H. (2004). Budget deficits and the current account balance: New evidence from panel data. *Journal of Economics and Finance*, *28*(1), 39–45. https://doi.org/10.1007/BF02761453

Nikiforos, M., Carvalho, L., & Schoder, C. (2015). "Twin deficits" in Greece: In search of causality. *Journal of Post Keynesian Economics*, *38*(2), 302–330. https://doi.org/10.1080/01603477.2015.1065675

Obadic, A., Globan, T., & Nadoveza, O. (2014). Contradicting the twin deficits hypothesis: The role of tax revenues composition. *Panoeconomicus*, *61*(6), 653–667. https://doi.org/10.2298/PAN1406653O

Ogbonna, B. (2013). Twin deficits or Ricardian equivalence hypothesis: evidence from Nigeria. *Journal of Banking*, *7*(1), 1–48.

Piersanti, G. (2000). Current account dynamics and expected future budget deficits: Some international evidence. *Journal of International Money and Finance*, *19*(2), 255–271. https://doi.org/10.1016/S0261-5606(00)00004-8

Salvatore, D. (2006). Twin deficits in the G-7 countries and global structural imbalances. *Journal of Policy Modeling*, *28*(6), 701–712. https://doi.org/10.1016/j.jpolmod.2006.06.003

Tosun, M. U., Iyidogan, P. V., & Telatar, E. (2014). The Twin Deficits in Selected Central and Eastern European Economies: Bounds Testing Approach with Causality Analysis. *Romanian Journal for Economic Forecasting*, *17*(2), 141–160.

Trachanas, E., & Katrakilidis, C. (2013). The dynamic linkages of fiscal and current account deficits: New evidence from five highly indebted European countries accounting for regime shifts and asymmetries. *Economic Modelling*, *31*, 502–510. https://doi.org/10.1016/j.econmod.2012.12.026

Turan, T., & Karakas, M. (2018). Asymmetries in twin deficit hypothesis: evidence from CEE countries. *Ekonomický Časopis (Journal of Economics)*, *66*(6), 580–597.

Vamvoukas, G. A. (1999). The twin deficits phenomenon: Evidence from Greece. *Applied Economics*, *31*(9), 1093–1100. https://doi.org/10.1080/000368499323571

Zildžović, E. (2015). The sustainability of Serbia's external position: The impact of fiscal adjustment and external shocks. *Economic Annals*, *60*(204), 31–60. https://doi.org/ 10.2298/EKA1504031Z

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Branislav Boričić[*]

A NOTE ON DICTATORSHIP, LIBERALISM AND THE PARETO RULE

ABSTRACT: *This note presents a short comment and explanation of the approach and results presented in my previous papers published in this journal. I have divided the traditional social choice axioms, introduced by Kenneth Arrow and Amartya Sen, into two classes, based on their linguistic and mathematical complexity. The first class consists of 'the unrestricted domain' and 'the independence of irrelevant alternatives', (Arrow, 1963; (Sen, 1970b;Maskin, 2020), which need a higher-order language, and can be treated as meta-axioms. The second class contains a group of linguistically simpler axioms, such as 'dictatorship', 'liberalism' and 'the Pareto rule'. Naturally, it is possible to make an easier logical analysis of the deductive properties and relationships between the axioms belonging to the second class, and the paper explains a method for their simplification. The basic conclusion is that after these simplifications, we obtain a fragment of the traditional Arrow-Sen theory in which we can also prove well-known impossibilities, including the counterparts of Arrow's and Sen's theorems. I consider that the value of each simplified approach lies in providing an opportunity to a wider circle of readers to better understand the basic ideas, results and spirit of traditional Social Choice Theory.*

KEY WORDS: *impossibility theorems; social choice theory; dictatorship; liberalism; Pareto rule.*

JEL CLASSIFICATION: D72, D71.

AMS Mathematics Subject Classification (2020): 91B14, 03B10, 91B02.

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Elements of traditional Social Choice Theory relevant to this note are given in Fishburn (1973) and Sen (1970a, 1970b, 1995), and a deep analysis of the language and logic of this theory was presented in Routhley (1979).

The results presented in Boričić (2009, 2014a, 2014b) cannot be considered as a repetition of the well–known theorems of Social Choice Theory, but as their analogues in a new context, subtly simplified and modified, but similar. Here, we want to explain the methodological and logical background of axiomatizations of fragments of traditional Social Choice Theory presented in our previous works Boričić (2009, 2014a, 2014b) or Boričić, & Srećković (2023). Namely, in both papers Boričić (2009, 2014b) the Pareto rule, and the dictatorship and liberalism axioms are given in an essentially different, but very similar and recognizable form. During our research, it was intuitively clear that we obtained some simpler fragments of traditional Social Choice Theory, but their pure logical relationships and methodological argumentations were absent. Now, in hindsight we are able to formally explain the status of these simplified forms of traditional axioms.

The traditional axioms employ quantification over relations and combine natural and higher–order formal languages, such as follows:

The Pareto rule TP, prefixed by **T** to denote a traditional form of **P**, claims that, for **all profiles** $\mathcal{P}$ and all alternatives $x, y \in X$, if every individual $i \in V$ prefers x to y, then society must prefer x to y. This is, in fact, a weak version of the Pareto principle, as introduced by Kenneth Arrow (see Arrow (1963) or Sen (1970a, 1970b, 1995)).

The dictatorship axiom TD states that there is a person $i \in V$, a dictator, having such power that, for **all profiles** $\mathcal{P}$ and all alternatives $x, y \in X$, if i prefers x to y, then society must prefer x to y as well (see Arrow (1963) or Sen (1970a, 1970b)).

The liberalism axiom TL supposes that, for **all profiles** $\mathcal{P}$ and each individual $i \in V$ there is at least one pair of alternatives $(x, y) \in X^2$ such that $x \neq y \wedge (xP_iy \to xPy) \wedge (yP_ix \to yPx)$ (see Sen (1970a, 1970b)).

These conditions can be respectively presented more formally in the following way:

$$\mathbf{TP}: \quad (\forall \mathcal{P})(\forall x, y \in X)((\forall i \in V)xP_iy \to xPy),$$

$$\mathbf{TD}: \quad (\exists i \in V)(\forall \mathcal{P})(\forall x, y \in X)(xP_iy \to xPy)$$

and

$$\mathbf{TL}: \quad (\forall\mathcal{P})(\forall i \in V)(\exists x, y \in X)(x \neq y \wedge (xP_i y \to xPy) \wedge (yP_i x \to yPx))$$

Meanwhile, in Boričić (2009, 2014b), we use their simplified variations such as:

$$\mathbf{SP}: \quad (\forall x, y \in X)((\forall i \in V)xP_i y \to xPy),$$

prefixed by **S** to denote a simplified form of **P**,

$$\mathbf{SD}: \quad (\exists i \in V)(\forall x, y \in X)(xP_i y \to xPy)$$

and

$$\mathbf{SL}: (\forall i \in V)(\exists x, y \in X)(x \neq y \wedge (xP_i y \to xPy) \wedge (yP_i x \to yPx))$$

supposing that these variations hold for **all profiles** $\mathcal{P}$, which is in line with the general assumption about the schematic character of axioms.

If we employ the deduction relation $A \vdash B$ to denote that "B can be derived from A", as in Boričić (2009, 2014b), we note that in all cases we have:

$$\mathbf{TP} \vdash \mathbf{SP}, \quad \mathbf{TD} \vdash \mathbf{SD} \text{ and } \mathbf{TL} \vdash \mathbf{SL}$$

i.e. that each simplified form is deductively entailed by an appropriate traditional form, fact which is based on the following general logical law:

$$\exists x \forall y A \vdash \forall y \exists x A$$

and then moving the universal quantification $\forall y$ to some kind of metatheoretical level. This operation can be of great importance when the object '$\forall y$' belongs essentially to a higher–order language, such as $\forall\mathcal{P}$. By this procedure we can obtain a similar but essentially simpler fragment of the theory which could be more approachable than the original one. Analogue statements were expressed in the first–order language, in Boričić (2009) and Boričić, & Srećković (2023), and in an almost propositional language, in Boričić (2014a, 2014b).

Finally, let us consider an example concerning famous Arrow's impossibility theorem. This theorem can be formulated as **TP** $\vdash$ **TD**, assuming that conditions of 'unrestricted domain' and 'the independence of irrelevant alternatives' hold, in original Arrow's theory, while its analogue, a similar statement, in this new simplified context, is the following one: $\mathbf{SP} \vdash \mathbf{SD}$. Let us emphasize that neither counterpart $\mathbf{SP} \vdash \mathbf{SD}$ implies Arrow's original theorem $\mathbf{TP} \vdash \mathbf{TD}$, nor vice versa. Consequently, these two statements can be considered as two roughly connected facts in two parallel worlds. Similarly, we can present a counterpart of Chichilnisky's original theorem Chichilnisky (1982), 'the impossibility of a non–Paretian dictator': $\mathbf{TD} \vdash \mathbf{TP}$, and its counterpart in our simplified context: $\mathbf{SD} \vdash \mathbf{SP}$, asserting again that there is no immediate formal logical connection between these two statements. But, on the other side, bearing in mind that $\mathbf{TP} \vdash \mathbf{SP}$ and $\mathbf{TL} \vdash \mathbf{SL}$, we can directly derive well–known Sen's 'impossibility of a Paretian liberal': $\mathbf{TP}, \mathbf{TL} \vdash$, from its simplified version $\mathbf{SP}, \mathbf{SL} \vdash$, meaning that the axioms $\mathbf{SP}$ and $\mathbf{SL}$, and, consequently, the axioms $\mathbf{TP}$ and $\mathbf{TL}$, when they appear together, make the theory inconsistent.

I consider that the value of this simplified approach is in giving an opportunity to a wider circle of readers to better understand the basic ideas and results of traditional Social Choice Theory.

REFERENCES

Arrow, K. J. (1963). *Social Choice and Individual Values,* John Wiley, New York.

Boričić, B. (2009). Dictatorship, liberalism and the Pareto rule: possible and impossible, *Economic Annals*, LIV 181, 45–54.

Boričić, B. (2014a). Multiple von Wrights preference logic and social choice theory, *The Bulletin of Symbolic Logic,* 20, 224–225.

Boričić, B. (2014b). Impossibility theorems in multiple von Wright's preference logic, *Economic Annals*, LIX 201, 69–84.

Boričić, B., & Srećković, M. (2023). Vetoing – social, logical and mathematical aspects, (to appear)

Chichilnisky, G. (1982). The topological equivalence of the Pareto condition and the existence of a dictator, *Journal of Mathematical Economics,* 9 223–233.

Fishburn, P. C. (1973). *The Theory of Social Choice,* Princeton University

Press, Princeton.

Maskin, E. (2020). A modified version of Arrow's **IIA** condition, *Social Choice and Welfare*, 54, 203–209.

Routhley, R. (1979). Repairing proofs of Arrow's general impossibility theorem and enlarging the scope of the theorem, *Notre Dame Journal of Formal Logic,* 20, 879–890.

Sen, A. K. (1970a). The impossibility of a Paretian liberal, *Journal of Political Economy,* 78 (1), 152–157.

Sen, A. K. (1970b). *Collective Choice and Social Welfare,* Holden–Day, San Francisco, (Fourth edition: Elsevier, Amsterdam, 1995).

Sen, A. K. (1995). Rationality and social choice, *American Economic Review,* 85, 1–24.

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ERRATUM

CORRECTION TO ARANDARENKO, PAVLOVIĆ (2023) A correction has been made to the article "EGALITARIANISM AND REDISTRIBUTIVE REFORM IN SERBIA AFTER 2000" by Mihail Arandarenko and Dušan Pavlović in Economic Annals, 2023, LXVIII (237): 7- 36

On the journal contents page there was an error: the order of the authors is written incorrectly "Dušan Pavlović, Mihail Arandarenko" should be: "Mihail Arandarenko, Dušan Pavlović…". The correct order of authors is written in the paper itself (https://doi.org/10.2298/EKA2337007P).

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Economic Annals is an international professional journal published quarterly by the Faculty of Economics and Business, University of Belgrade. The journal publishes research in all areas of economics and business. It publishes high-quality research articles of both theoretical and empirical character. The journal especially welcomes contributions that explore economic issues in comparative perspective with a focus on Southeast Europe and the wider European neighbourhood. Any paper submitted to the *Economic Annals* should **NOT** be under consideration for publication by other journals or publications. **Contribution written in English should be submitted electronically to ScholarOne**.

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Rodrik, R., Subramanian, D., & Trebbi, F. (2004). Institutions rule: the primacy of institutions over geography and integration in economic development. *Journal of Economic Growth*, 9(2), 131-165.

https://DOI: 10.1023/B:JOEG.0000031425.72248.85.

• Books

Author surname, initial(s). (Year). *Title*. Publisher location: Publisher

De Grauwe, P. (2020) *Economics of Monetary Union* (13th ed.). Oxford: Oxford University Press.

• Edited Book

Author surname, initial(s). (Ed(s).). (Year). *Title*. Publisher location: Publisher

Baltagi, B.H. (Ed.). (2003). *A Companion to Theoretical Econometrics*. Oxford: Blackwell

• Book with several authors

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Baumol, W. J., Panzar, J. C., & Willig, R.W. (1982). *Contestable Markets and the Theory of Industry Structure*. New York: Harcourt, Brace, Jovanovich, Inc.

• Chapter in Book

Last name of the chapter author, initial(s). (Year). Chapter title. In editor initial(s), surname (Ed.). *Title* (ed., pp.). Publisher location: Publisher

McMillan J., & Woodruff C. (2003) The central role of entrepreneurs in transition economies. In G. S. Fields, & G. Pfefferman (Eds.). *Pathways Out of Poverty* (pp. 105-121). Dordrecht: Springer. https://doi.org/10.1007/978-94-010-0009-3_6.

• E-Book

Author surname, initial(s). (Year). *Title*. URL

Perry, R.B. (1909). *The Moral Economy.*

https://manybooks.net/book/137844/read#epubcfi(/6/2[id00000]!/4/2[id00000]/1:0)

• Technical Reports or Working Papers

Individual authors

Author surname, initial(s) or corporate name. (Year). *Title*. (Report or Working Paper No.). URL.

Cătuți, M., Kustova, I. and Egenhofer, C. (2020) *Delivering the European Green Deal for Southeast Europe: Do we need a regional approach?* (CEPS Research Report No.2020/1). https://www.ceps.eu/wp-content/uploads/2020/06/RR_2020-01_European-Green-Deal-for-South-Eastern-Europe.pdf.

Corporate authors

American Psychological Association, Task Force on the Interface Between Psychology and Global Climate Change. (2009). *Report of the APA Task Force on the Interface Between Psychology and Global Climate Change.*

http://www.apa.org/science/about/publications/climate-change.aspx

• Newspaper Articles

Author surname, initial(s). (Year, Month Day). Title. *Title of Newspaper*, p. or pp. URL*

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Note: the date includes the year, month and date.

Smialek, J. (2020, May 2). Hotel Group Will Return Tens of Millions in Small Business Loans. *The New York Times*, pp. 10.

https://www.nytimes.com/2020/05/02/business/economy/ashford-hotels-virus-monty-bennett.html

• Website

Author surname, initial(s). (Year, month day). *Title*. URL

Mitchell, J.A. (2017, May 21). *How and when to reference*.

https://www.howandwhentoreference.com