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The response of inflation to budgetary shocks in Russia: an (S)VAR approach in economically uncertain times

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Aim: The motivation for this research stems from Russia's notable high levels of government expenditure due to its continual involvement in armed conflicts, which often result in budgetary imbalances and economic policy uncertainty. These factors impact the inflation rate. This study delves into the complex relationship between budgetary shocks, economic uncertainty, and inflation within the context of the Russian economy. The aim of this study is to unravel how budgetary decisions, made amidst a globally uncertain economic environment, influence inflation dynamics. In other words, the objective of this study is to analyse the effects of budgetary shocks on the inflation rate in Russia, taking into account the uncertain context of her economic policy.

Design / Research methods: To achieve the objective, we employ a Structural Vector AutoRegression (S)VAR approach, covering the period from 2003-Q1 to 2022-Q4. This methodological approach allows for a comprehensive analysis of how economic uncertainty influences the identification of budgetary shocks within an (S)VAR model.

Conclusions / findings: The findings underscore that incorporating the economic uncertainty index into the model yields statistically significant estimates, suggesting that variations in economic uncertainty shape the relationship between budgetary shocks and inflation. This sheds light on the intricate mechanisms through which economic uncertainty influences the behaviours of economic agents and policy decisions, thereby affecting the transmission of budgetary shocks to inflation. In contrast, without the economic uncertainty index, the response of the inflation rate to budgetary shocks is insignificant.

Originality / value of the article: This study makes an original contribution by incorporating the Economic Uncertainty Index to better capture budgetary shocks. By showing how uncertainty affects the effectiveness of fiscal policies on inflation, it offers new perspectives on macroeconomic stability. This

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approach provides a more detailed analysis of the responses of economic actors and their implications for policymakers in volatile economic environments.

Keywords: Russia, economic policy uncertainty, inflation, budgetary shocks, (S)VAR.
JEL: E31; E62; C51

1. Introduction

Russia's economic resilience stems from its dependency on oil and gas exports, which contribute to 30% and 45% of the country's total exports, respectively, according to the Russian Ministry of Finance. This structural characteristic of its economy becomes particularly sensitive during periods of instability due to armed conflicts that often lead to increased government expenditures, triggering consequential adjustments in fiscal policy. This, in turn, directly affects the economic equilibrium by influencing various dimensions of the economy.

Indeed, such a course of action results in an increase in military expenditures, subsequently causing a rise in government spending and a decline in tax revenues (Itskhoki, Mukhin 2022). For instance, since the onset of the conflict in Ukraine in February 2022, government expenditures within the Russian Federation has surged by 25% between the first and third quarters, while tax revenues experienced a 20% decrease during the same period. These budgetary fluctuations are accompanied by a rise in the inflation rate, which escalated to 21.6% in November 2022, in contrast to the 6.7% observed in November 2021 (Russian Ministry of Finance). Given the pivotal role inflation plays in transmitting effects among macroeconomic aggregates (Diop, Diaw 2015), such circumstances have prompted significant interest in empirical research concerning the impact of fiscal shocks on the inflation rate.

In this context, we aim to examine the repercussions of fiscal shocks on inflation in Russia by adopting a multivariate analytical approach using the structural vector autoregressive (S)VAR model.

Analysing the effects of fiscal policy shocks within the framework of a (S)VAR model remains challenging due to the difficulty in identifying structural shocks (Ramey 2011). Since the emergence of the neo-Keynesian model, researchers have explored various indicators to capture structural shocks in a timely manner. Among these indicators are inflation predictions and the output gap, which offer informed perspectives on the economic situation by revealing business cycles. However, it is essential to note that these indices, while valuable, remain unobservable and involve monetary aggregates such as interest rates, potentially leading to less robust estimations when introduced into the model (Mazzi et al. 2016).

According to Keynes (1936: 186), the uncertainty of economic policy leads economic actors to adopt a cautious approach towards consumption and investment. This uncertainty is evaluated using three fundamental elements (Baker et al. 2014): fiscal provisions, forecast divergences from consensus, and the number of articles published in official journals incorporating terms like "uncertain" or "economic uncertainty," along with other relevant terms related to economic policies.[1] Consequently, we propose that considering economic uncertainty in Russia can help us understand the stance of its fiscal policy and can provide the (S)VAR model with the means to determine structural shocks. Our study therefore aims to examine the response of the inflation rate to fiscal shocks in Russia over the period 2003Q1–2022Q4 within a context of economic policy uncertainty. The innovative contribution of our study lies in integrating economic uncertainty, measured by the Economic Policy Uncertainty Index, as a determining factor in analysing fiscal shocks and their impact on inflation. This approach introduces a novel factor that helps capture shocks within the (S)VAR model. By exploring the interaction between economic uncertainty and fiscal policies, our research provides a new and essential dimension to the understanding of Russian economic dynamics, a perspective that has not yet been fully explored in the existing literature.

We formulate two main hypotheses: 1) A positive budgetary shock in an uncertain economic environment results in a significant increase in the inflation rate; 2) A negative budgetary shock in an uncertain economic environment leads to a decrease in the inflation rate.

The rest of the paper is organised as follows: section two provides a brief overview of the existing literature focusing on the impact of fiscal shocks on

[1] This is the standard deviation among economists' forecasts for the variables (inflation and government spending), which serve as measures of economic policy uncertainty. The economists are selected by the Philadelphia Fed and the comprehensive list is available at: https://www.philadelphiafed.org/research-and-data/economists (BSI Economics, Quentin Blanc 2015).

macroeconomic aggregates, with a specific emphasis on the inflation rate. Section three elaborates on the methodological framework of the (S)VAR model. Section four outlines the outcomes extracted from our analysis. Section five provides an interpretation of the findings based on pertinent economic theory. Lastly, Section 6 summarises our concluding remarks.

2. Research background

The topic of interactions between fiscal shocks and macroeconomic aggregates, particularly the inflation rate, has often sparked conflicting theoretical and empirical debates.

On a theoretical level, the post-Keynesian theory provides a valuable framework for analysing the effects of fiscal shocks on inflation in the context of economic uncertainty. Under this perspective, a notable effect is that government spending can play a stabilising role in times of economic uncertainty by supporting overall demand without necessarily generating excessive inflation. This phenomenon is explained by the fact that, in conditions of uncertainty, economic actors and consumers tend to save rather than spend, which alleviates inflationary pressures (Davidson 1991).

On an empirical level, the analysis of fiscal shocks began with pioneering work by Blanchard and Perotti (2002), who used a structural vector autoregressive (S)VAR model to study the post-war American economy. Their goal was to investigate the effects of shocks related to government expenditures and fiscal constraints on economic growth and private consumption. Their findings indicated that positive expenditure shocks led to increases in both growth and consumption, suggesting that increased government expenditures can stimulate economic activity by boosting demand. Conversely, positive tax shocks resulted in decreases, implying that higher taxes reduce disposable income and thus lower consumption and economic activity.

Perotti (2004) used a VAR(S) model to examine the impact of fiscal shocks on inflation in five OECD countries from 1961 to 2001. The results show that fiscal shocks often have short-lived effects on inflation. This suggests that fiscal policy may not be a good way to change the economy over the long term. This is likely because shocks are not captured well enough in this model.

Mountford (2005) employed an (S)VAR model with American data to show that tax revenue shocks negatively impact private consumption and gross domestic product (GDP). Higher taxes reduce consumer spending and overall economic output. In contrast, government expenditure shocks did not reduce consumption but instead created a crowding-out effect on private investment, where increased government expenditures lead to reduced private sector investment, ultimately lowering the inflation rate by reducing aggregate demand.

Afonso and Sousa (2009) applied a Bayesian (S)VAR approach to the United States, Germany, Italy, and the United Kingdom. They found that government expenditure shocks did not affect inflation, indicating that government expenditure increases did not translate into higher prices.

Burriel et al. (2010) used a (S)VAR model for the euro area and found that the inflation response to fiscal shocks was neutral. This neutrality implies that fiscal shocks have not consistently resulted in changes in the inflation rate. However, the output response depended on the budget stress index. When the budgetary tension index was high, inflation tended to rise as a result of positive shocks in government spending, indicating that fiscal uncertainty exacerbates inflationary pressures. On the contrary, when fiscal uncertainty was low, inflation tended to decline, suggesting that fiscal insecurity mitigated the inflationary impact of government spending. According Davidson (1991), the budget stress index has increased economic uncertainty by fuelling uncertainties about financial stability, which diminishes the confidence of economic players.

Baum and Koester (2011) conduct an analysis using a Threshold VAR model and find that the impact of fiscal shocks on economic activity and inflation depends heavily on the initial level of economic uncertainty. They note that high levels of uncertainty can weaken the economy’s response to fiscal shocks, often by reducing the expected stimulus effect on growth and mitigating inflationary pressures.

Auerbach and Gorodnichenko (2012), show in their study that the effectiveness of fiscal policies varies significantly depending on the level of economic uncertainty. In periods of high uncertainty, the effects of fiscal shocks are less predictable and may have differentiated impacts on inflation. This suggests that economic uncertainty alters the transmission of fiscal policies to the real economy.

The Ramey and Zubairy study (2014) analyses the multipliers of government spending in the United States based on the uncertain economic context. They found that the effects of government expenditure on economic growth vary significantly during periods of low economic uncertainty compared to periods with high economic insecurity, where spending effectiveness may be affected by the level of increased economic insecurity.

Beetsma et *al*. (2015), their study deals with the effects of fiscal consolidation on the confidence of economic operators in Europe. They find that fiscal consolidation measures can hurt confidence, thereby exacerbating economic uncertainty and potentially curbing economic recovery by increasing inflation and decreasing consumption.

Fort et *al*. (2017), although his paper does not deal directly with fiscal shocks, it examines employment trends in the manufacturing sector in the United States. It highlights how economic uncertainty can influence employment dynamics through government spending.

Olamid et *al*. (2022) use the (S)VAR method to analyse the impact of fiscal and monetary shocks on the economic dynamics of the East African Community (CAE). By focusing on the effects on the inflation rate, exchange rate, and GDP of member countries, it aims to understand how these variables interact in a context of persistent economic uncertainty.

In summary, this literature review highlights that despite the recognition of a significant link between fiscal shocks and inflation in the literature, there is a noticeable variability in the importance of economic uncertainty as a moderator of this relationship. Some works pay particular attention to this factor, while others treat it in a more marginal way. This diversity of perspectives underlines the importance of our study to deepen these nuances and strengthen the reliability of our findings in this complex area of economic research.

3. Methodology

This section elucidates the methodological approach of the (S)VAR model along with the selected data, aiming to address the objective of the present study.

3.1.The theoretical framework of the (S)VAR approach

Adopting the (S)VAR approach as presented by Blanchard and Perotti (2002: 1329–1368) offers several advantages: *i*) it relies on economic theory to interpret shocks, establishing conceptual links between variables in contrast to a standard VAR, which generates economically uninterpretable innovations (D'amico, King 2023); *ii*) it simplifies the identification of policy-related shocks by employing time lags to determine when a shock occurs and when corrective measures are enacted (Perotti, 2002); *iii*) it accounts for the simultaneity of effects among variables, an essential feature when analyzing interactions between variables (Kuma 2018); and *iv*) it provides insights into the characteristics of fiscal policy instruments in relation to economic activity (Burriel et *al*. 2010).

Our starting point (Step 1) is the primitive form of the VAR model, formulated as follows:

$$AY_t = \lambda + \sum_{i=1}^{P} \beta_i Y_{t-i} + v_t \tag{1}$$

A: square matrix of order (k×k); Y_t: vector of k endogenous random variables of order (k×1); $\boldsymbol{\lambda}$**:** constant terms vector; *P:* lag order; β_i: Square matrix of coefficients associated with lags of order (k×k); v_t**:** diagonal matrix of error terms for the system of order (k×1).[2]

We assume weak stationarity for input variables, as well as an absence of autocorrelation among the error terms of equation (1) (Ljung-Box P-value > 0.05) for each lag. Additionally, a normal distribution of error terms is recommended (Jarque-Bera P-value > 0.05) to validate the VAR model.

By validating the VAR model, we proceed to the configuration of the (S)VAR for its estimation (Step 2), as follows:

[2] The assumption of matrix diagonalization $v_t \sim iid(0,\sigma_v^2)$ highlights the possibility of defining orthogonal structural shocks in pairs, which are at the core of the main objective of the (S)VAR.

$$Y_t = A^{-1}\lambda + \sum_{i=1}^{p} A^{-1}\beta_i Y_{t-i} + A^{-1}v_t \qquad (2)$$

$A^{-1}\lambda = \boldsymbol{\alpha}$; $A^{-1}\beta_i = \boldsymbol{\phi_i}$; $A^{-1}v_t = \boldsymbol{e_t}$

Where,

β_i: Diagonal matrix with non-zero elements only on its main diagonal; A^{-1}: The inverse of the unit triangular lower matrix A. We assume that the responses to shocks are likely to be observed in the own lagged values of each variable, thus the matrix A is lower diagonal (Kuma 2018).The equation (2) in its reduced form is therefore:

$$Y_t = \alpha + \sum_{i=1}^{p} \phi_i Y_{t-i} + e_t \qquad (3)$$

The matrix A of order ($k \times k$) in its primitive form (equation 1) represents an identity matrix, employed in constructing the Structural Vector Autoregressive (S) VAR model to facilitate shock identification. Equation (2) $v_t = Ae_t$, elucidates the relationship between unobserved shocks and endogenous variables. e_t Signifies unobserved disturbances impacting endogenous variables, with its economic significance discernible only through its linear combination with instantaneous structural shocks (Diop, Diaw 2015).

The autoregressive elasticities of matrix A mirror the influence of past values of variables on their current values. They illustrate how each endogenous variable depends on its own past values. However, they are not directly linked to shocks. The elasticities of matrix B quantify the propagation of these shocks to the observed variables.

3.2.Data and source

The variable selection for our (S) VAR model applied to the Russian economy is based on the theoretical framework developed by Burriel et *al*. (2010) and the foundational theoretical rule.

The vector Y_t is, thus, composed of ($p_t, y_t, r_t, c_t, u_t, t_t, g_t$), where:

- (p) Inflation rate is represented by the Consumer Price Index (CPI) (excerpted from Concise Economic Indicators Compilation Flex database).
- (c)Private final consumption is expressed in the Russian national currency, the ruble (excerpted from Global Economy database).

- (*r*) Short-term interest rate measuring the yield for 3 months up to 90 days, expressed as a percentage (excerpted from Federal Reserve Economic Database). [3]
- (*u*) Index of economic policy uncertainty expressed in points (excerpted from Federal Reserve Economic Database).[4]
- (*y*) Production is represented by the real Gross Domestic Product (GDP), expressed in the Russian national currency, the ruble (excerpted from Federal Reserve Economic Database).
- (*t*)Total tax revenues expressed as a percentage of GDP (excerpted from Global Economy database).[5]
- (*g*) Government consumption expenditures expressed in the Russian national currency, the ruble (excerpted from Concise Economic Indicators Compilation Flex database).[6]

Except for the variables (*u*), (*y*) and (*t*), the remaining variables are measured at constant prices (100 = 2015). The data are transformed into logarithms. Data have been collected at a quarterly frequency, spanning from 2003:Q1 to 2022:Q4.[7]

3.3.Theoretical identification of short- term structural constraints

Given our interest in analysing the inflation response to budgetary shocks, it becomes imperative to examine the linear combination formed by inflation residuals, expenditure residuals, and revenue tax residuals within the simplified formulation of the system (Burriel, 2010). Furthermore, since our modelling approach is situated

[3]The selection of a short-term interest rate relates to its connection with the inflation rate, which impacts private consumption and investment decisions (Patterson, Lygnerud 1999).

[4] The data were collected at a monthly frequency, and we computed the arithmetic mean over three months corresponding to each quarter.

[5]We opt for the utilisation of total tax revenues as they encompass income derived from the value-added tax (VAT) as well as direct taxation, such as the Global Income Tax (GIT), both of which exert an influence on the inflation rate due to the portion of savings derived from individual incomes (Mountford 2005).

[6]We have chosen to consider government consumption expenditures for several reasons: our study highlights military expenditures, which are accounted for within the consumption category in terms of national accounting; Consumption expenditures also encompass the repayment of public debt;Contemporary literature places particular emphasis on consumption expenditures due to their significant impact on aggregate demand within the economy (Keynes 1936).

[7]The chosen period encompasses events marked by instability across various dimensions: financial, economic, health-related, and political.

within a structural context, it is crucial to present the linear combination of structural shocks from other variables, contributing to the determination of the matrix of structural shocks.

To establish the constraints, we draw upon Perotti's (2004) approach, which suggests that quarterly variables respond to structural shocks with a lag exceeding three (3) months, implying that the instantaneous responses of certain variables to shocks are null.

The number of constraints required in an (S)VAR model is contingent upon the research objectives as well as the complexity of inter-variable relationships.[8] Furthermore, it is important to note the pivotal role that the order of variables plays in setting up constraints. The system of structural equations within the framework of our (S)VAR model is as follows:

$$
\begin{aligned}
v_t^p &= \phi_{p,c} v_t^c + \phi_{p,r} v_t^r + \phi_{p,u} v_t^u + \phi_{p,y} v_t^y + \beta_{p,t} e_t^t + \beta_{p,g} e_t^g + e_t^p \\
v_t^c &= \phi_{c,p} v_t^p + \phi_{c,r} v_t^r + \phi_{c,u} v_t^u + \phi_{c,y} v_t^y + \beta_{c,t} e_t^t + \beta_{c,g} e_t^g + e_t^c \\
v_t^r &= \phi_{r,p} v_t^p + \phi_{r,c} v_t^c + \phi_{r,u} v_t^u + \phi_{r,y} v_t^y + \beta_{r,t} e_t^t + \beta_{r,g} e_t^g + e_t^r \\
v_t^u &= \phi_{u,p} v_t^p + \phi_{u,c} v_t^c + \phi_{u,r} v_t^r + \phi_{u,y} v_t^y + \beta_{u,t} e_t^t + \beta_{u,g} e_t^g + e_t^u \\
v_t^y &= \phi_{y,p} v_t^p + \phi_{y,c} v_t^c + \phi_{y,r} v_t^r + \phi_{y,u} v_t^u + \beta_{y,t} e_t^t + \beta_{y,g} e_t^g + e_t^y \\
v_t^t &= \phi_{t,p} v_t^p + \phi_{t,c} v_t^c + \phi_{t,r} v_t^r + \phi_{t,u} v_t^u + \phi_{t,y} v_t^y + \beta_{t,g} e_t^g + e_t^t \\
v_t^g &= \phi_{g,p} v_t^p + \phi_{g,c} v_t^c + \phi_{g,r} v_t^r + \phi_{g,u} v_t^u + \phi_{g,y} v_t^y + \beta_{g,t} e_t^t + e_t^g
\end{aligned}
$$

e_t^i,$v_t^i$are the structural shocks, and the residual terms of the variables, respectively.

Our initial approach prioritises variables subject to constraints (Perotti 2004) to ensure that the matrix structure (A) remains coherent and to promptly introduce relationships mandated by the constraints.

Following the reasoning advocated by Burriel (2010), we formulate short-term restrictions as follows:

[8]The number of restrictions to be retained in the (S)VAR is given by: n = k(k-1)/2; where, k represents the number of endogenous variables (Blanchard, Perotti 2002). Nevertheless, it is imperative to refer to economic definitions and the outcomes of non-causality tests in various previous studies. The number of restrictions for our model is n = 21.

– An instantaneous shock to the inflation rate has no effect on private consumption (*c*), interest rate (*r*), economic uncertainty (*u*), production (*y*), tax revenues (*t*), or government expenditures (*g*), which entails(6) restrictions:

$$\phi_{p,c} = \phi_{p,r} = \phi_{p,u} = \phi_{p,y} = \beta_{p,t} = \beta_{p,g} = 0.$$

– An instantaneous shock to private consumption (*c*) does not influence economic uncertainty (*u*), production (*y*), tax revenues (t), or government expenditures (*g*) which entails (5) restrictions:

$$\phi_{y,r} = \phi_{y,u} = \phi_{y,y} = \beta_{y,t} = \beta_{y,g} = 0.$$

– An instantaneous shock to the interest rate (*r*) has no effect on economic uncertainty (*u*), production (*y*), tax revenues (*t*), or government expenditures (*g*) which entails(4) restrictions:

$$\phi_{r,c} = \phi_{r,u} = \phi_{r,y} = \beta_{r,t} = \beta_{r,g} = 0.$$

– An instantaneous shock to economic uncertainty (*u*) has no effect on production (y), tax revenues (*t*), or government expenditures (*g*) which entails (3) restrictions:

$$\phi_{u,y} = \beta_{u,t} = \beta_{u,g} = 0$$

– An instantaneous shock to production (*y*) does not react to tax revenues (*t*) and government expenditures (*g*) which entails (2) restrictions:

$$\beta_{y,t} = \beta_{y,g} = 0.$$

– An instantaneous shock to tax revenues (*t*) does not react to government expenditures (*g*) which entails (1) restriction:

$$\beta_{t,g} = 0.$$

4. Results

Table 1 provides the statistical description of the variables utilised.

Table 1.Statistical description of the data

	l(*p*)	l(*y*)	l(*r*)	l(*c*)	l(*u*)	l(*t*)	l(*g*)
Mean	4.319	16.818	2.036	4.551	5.077	2.645	4.296
Median	4.358	16.836	1.999	4.624	4.978	2.597	4.514
Maximum	5.053	17.132	3.051	5.079	6.559	3.163	5.195
Minimum	3.444	16.376	1.442	3.877	4.072	2.217	2.811
Std. Dev	0.452	0.156	0.334	0.246	0.611	0.236	0.677
Observation	80	80	80	80	80	80	80

Source: Authors'compilation.

4.1. (S)VAR model estimation

The initial step in estimating (S) VAR involves examining the stationarity of the variables (Table 2).

Table 2.Dickey-Fuller Augmented (ADF) and Phillips-Perron (PP) stationarity tests

		ADF			**'P**		
		I(0)	I (1)	I(2)	I(0)	I(1)	I(2)
l(p)	t- statistic prob	[-1.864] (0.663)	[-6.290]** (0.000)	-	[-1.838] (0.676)	[-6.306]** (0.000)	-
l(c)	t- statistic prob	[-2.241] (0.459)	[-5.686]** (0.000)	-	[-3.438] (0.053)	[-13.880]** (0.000)	-
l(r)	t- statistic prob	[-2.974] (0.146)	[-8.122]** (0.000)	-	[-2.960] (0.150)	[-8.122]** (0.000)	-
l(u)	t- statistic prob	[-6.086] (0.000)	-	-	[-6.090] (0.000)	-	-
l(y)	t- statistic prob	[-3.225] (0.088)	[-2.306] (0.422)	[-5.091]** (0.000)	[-3.189] (0.072)	[-2.012] (0.127)	[-6.619]** (0.000)
l(t)	t- statistic prob	[-0.884] (0.952)	[-7.580]** (0.000)	-	[-0.974] (0.941)	[-7.580]** (0.000)	-
l(g)	t- statistic prob	[-2.626] (0.270)	[-2.398] (0.397)	[-18.680]** (0.000)	[-2.826] (0.060)	[-9.044]** (0.000)	-

Source: Authors'compilation. ** The null hypothesis of non-stationarity is rejected at a 5% level.

Except for the economic uncertainty variable '*u*', all other variables are non-stationary at level I(0).

Subsequently, the optimal lag order (P) is determined using the Akaike Information Criterion (AIC) and Schwartz Criterion (SC). These criteria aim to select the model that provides the best data description, considering the limited sample size and the model's fit in terms of normality and absence of residual autocorrelation.[9]

The optimal lag order that minimises the AIC and SC criteria for our model is P=6.In the model estimated, the variables are significant (P-value < 0.05) at optimal lag: P=6. It is noteworthy at this juncture that the signs of the parameters are not extensively considered in the VAR model. Resolution stems from the estimation of the (S) VAR due to its incorporation of economic constraints (Perotti 2004).

4.2.Empirical identification of short-term structural constraints

By applying the conversion constraints to the residuals, which are formed by a linear combination of the orthogonal impulse response structure, we derive the matrix of contemporaneous relationships associated with the various instantaneous shocks as follows:

$$\begin{bmatrix} p \\ c \\ r \\ u \\ y \\ t \\ g \end{bmatrix} = \begin{pmatrix} 1.00 & 0.00 & 0.00 & 0.00 & 0.00 & 0.00 & 0.00 \\ 0.75 & 1.00 & 0.00 & 0.0 & 0.00 & 0.00 & 0.00 \\ -2.42 & 0.09 & 1.00 & 0.00 & 0.00 & 0.00 & 0.00 \\ 4.44 & 1.51 & 1.73 & 1.00 & 0.00 & 0.00 & 0.00 \\ -0.54 & -0.35 & 0.25 & -0.02 & 1.00 & 0.00 & 0.00 \\ 1.53 & -0.77 & 0.18 & -0.03 & -0.78 & 1.00 & 0.00 \\ 1.22 & -0.01 & 0.28 & -0.00 & -0.02 & 0.15 & 1.00 \end{pmatrix} \text{x} \begin{bmatrix} v_p \\ v_c \\ v_r \\ v_u \\ v_y \\ v_t \\ v_g \end{bmatrix}$$

The structural factorization of the matrix (*A*) reveals that variations in government consumption expenditures (*g*) and levels of economic uncertainty (*u*) exert a positive and significant impact on the inflation rate (+1.22; +4.44), which confirms the initial

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[9]The results of optimal lag selection and the validity testing of the model are provided in Appendix (A).

hypothesis. However, fluctuations in tax revenues (*t*) also lead to an increase in the inflation rate; thus, the second hypothesis of this study is refuted. These findings align with those of Sriyana (2019: 1701-1713) and Asandului et *al*. (2021: 899-919).

The variations in short-term interest rate (*r*) and production (*y*) demonstrate a negative and significant influence on the inflation rate (-2.42; -0.54). It is worth noting that the private consumption variable (*c*) does not exhibit statistical significance.[10]

4.3.Estimation of budgetary shocks

Our aim in this estimation is to provide an overview of the convergence between empirically determined budgetary shocks and the budgetary shocks that have occurred in the Russian economy (Figure 1).

The positive revenue shocks in Russia observed during the studied period are associated with *i*) administrative reform involving the introduction of new sections

in Part II of the Russian Federation Tax Code[11]; *ii*) incorporation of the Electronic Tax Register into savings banks to enhance control and reduce tax disputes; *iii*) the tax administration modernization project (2008); and *iv*) a substantial increase in the value-added tax rate (2019).

We also identify negative revenue shocks related to *i*) Federalism reforms and the reinforcement of fiscal system centralization; *ii*) Russia committing to abolish the interregional sales tax, leading to a significant revenue decrease estimated at nearly 5% for each region[12]; *iii*) social issues review resulting in a political crisis in January 2005 (Daucé, Walter 2006); *iv*) as part of economic revitalization through the National Projects, Russia reduced the taxation rate for self-employed workers ; and *v*) the

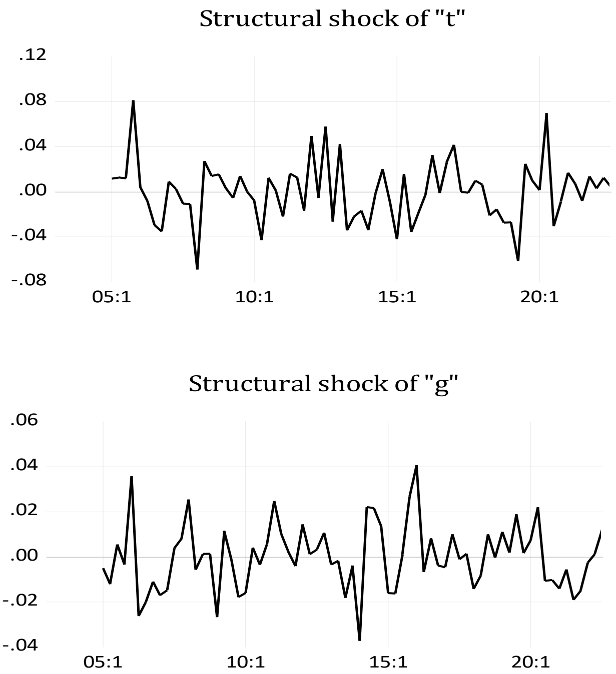
[10] The (S)VAR model estimation is presented in Appendix (B).

[11] During the years 2004, 2005, 2006, and 2007, the Federal Tax Service introduced new chapters to the Tax Code of the Russian Federation, namely Chapter 29: Tax on Gambling; Chapter 30: Social Goods; Chapter 25: Levies for the Use of Wildlife and Marine Biological Resources; Chapter 26: Unified Agricultural Tax – abolished sales tax; Chapters 25.3: State Duty and 31: Land Tax of the Tax Code of the Russian Federation; Water Tax, Chapter 02.25 of the Tax Code of the Russian Federation. Exemption from fees for the use of water bodies (Presidential Decree of the Russian Federation). Furthermore, the Ministry of Taxes and Levies of the Russian Federation underwent a transformation into the Federal Tax Service with the aim of enhancing the functions of executive bodies and optimizing tax management.

[12] Refer to Novikov (2005), where he discusses the reforms of federalist doctrine in Russia in his article.

COVID-19 health crisis led to reduced tax revenues due to decreased gross value added in the mining industry (Malkina 2021).

Figure1.Estimating budgetary shocks



Source: Authors'compilation.

Regarding positive government expenditure shocks, we observe: *i*) strong production growth triggering an expansionary budget policy aiming for significant public ownership expansion by the end of 2004 (Novikov 2005); *ii*) Russian military intervention in the Syrian conflict (2015); *iii*) preparations for a potential invasion of Ukrainian territory, illustrated by land, air, and maritime military manoeuvres, resulting in a significant increase in government expenditures (2021), and *iv*) Russian annexation of Ukrainian territories (2022).

As for negative government expenditure shocks, they are linked to economic regulation moments. *i*) Increased global demand in 2004 led to higher import rates, intensifying inflationary pressures. To control this inflation, Russian budgetary authorities adopted an approach of stringent government expenditure restriction; *ii*) the increase in foreign exchange reserves before the 2008 global financial crisis, driven by energy export tariffs, reached its peak level, enabling partial debt cancellation and a decrease in government expenditures. Moreover, there was: *iii*) an oil shock manifested by a drastic price drop of over 70% and *iv*) restructuring within the oil industry to reduce export duties and implicit fuel consumption subsidies (2019).

4.4.Structural variance decomposition

The analysis of the inflation rate's variance decomposition due to budgetary shocks is presented in Table 3.

According to the analysis of the variance decomposition, it is clearly observable that a budgetary shock can contribute up to 15.5% to the variation of the inflation rate. In comparison, a shock to tax revenues can contribute up to 4.4% to the variation of the inflation rate. We estimate that the contribution of budgetary shocks to inflation variability is relatively low.[13]

Table 3. Contribution of budgetary shocks to inflation variability

	T1	T2	T3	T4	T5	T6	T7	T8	T9	T10
Tax revenue shock	0,0	0,06	0,4	0,5	1,9	2,0	1,7	1,5	4,4	7,6
Government consumption expenditure shock	0,0	3,4	3,6	4,5	6,9	9,2	11,2	11,0	15,4	13,5

Source: Authors'compilation. 'T' stands for the quarter.

4.5.Inflation response to budgetary shocks in the absence of an economic policy uncertainty index

This section focuses on examining the effects of budgetary shocks on inflation after eliminating the economic policy uncertainty variable '*u*' and estimating the (S) VAR model, which comprises just six (06) variables.[14]

[13]The impulse responses to budgetary shocks are reported in Appendix (B).
[14]The (S)VAR model estimation in the absence of the variable (*u*) is presented in Appendix (C).

The estimation of the matrix (*A*) reveals that the autoregressive elasticities of the short-term interest rate and government consumption expenditure variables hold negative statistical significance (-2.91 [-4.12]; -0.72 [-3.08], respectively). Conversely, the coefficients of the other variables do not exhibit any statistical significance.

Despite the statistical insignificance of the autoregressive elasticity of tax revenues at optimal lag level, we persist in our analysis of structural shocks. We remove the constraints related to economic uncertainty (parameters in matrix *A* are set to 0) to continue our investigation.

The impulse response of the inflation rate to budgetary shocks does not hold any statistical significance. In other words, we are not able to adequately identify the effects of structural budgetary shocks due to a lack of information, notably related to the economic uncertainty variable and potentially other variables.[15]

5. Interpretation

Observing a modest positive inflation rate response to a positive shock in government consumption expenditures can be attributed to the fact that the rise in the overall price level does not stem from an increase in aggregate demand but rather from expenditures allocated to non-productive sectors such as defence. In this context, the accentuation of the persistence of government expenditure shocks in Russia is primarily associated with military expenditures. During economic uncertainty, economic agents are not motivated to enhance consumption levels when expenditures that stimulate aggregate demand remain at their current level. This response arises from economic agents' concern over economic uncertainty, which may influence their spending decisions, resulting in a mild inflation level increase.From a post-Keynesian perspective, such modest inflationary responses underscore the theory's emphasis on effective demand, the role of expectations and economic agents' behaviour, as well as the impact of institutional and public policy frameworks. The theory posits that

[15]The impulse responses to budgetary shocks in the (S)VAR model, in the absence of the variable (*u*), are reported in Appendix (C).

increases in government expenditures, particularly in non-productive sectors, during economic uncertainty may not significantly boost aggregate demand due to cautious consumer and business responses. This view suggests that economic actors adjust their behaviour based on expectations of future economic conditions, influencing the effectiveness of fiscal measures in stimulating the economy (Davidson 1991).

The explanation behind the modest positive inflation rate response to a positive shock in tax revenues lies in the following mechanism: when tax revenues increase through indirect taxes indexed to prices (such as VAT), the prices of consumer goods increase in value, leading to inflation growth (Sagramoso 2004). Conversely, revenues from income tax and social contributions remain constant in Russia, resulting in a favourable wage dynamic to compensate for the increased tax burden on consumer goods (Sokoloff 2005). This dynamic translates into a relatively mild inflation rate increase, aligning with the findings of Sagramoso (2004).

We consider the multipliers relatively low due to Russia's history of high inflation rates. Furthermore, the results exhibit a slight and statistically significant reduction in private consumption in response to tax revenue shocks, while the impact of consumption expenditure shocks holds no statistical significance. These findings are in line with the study by Jørgensen and Ravn (2022). Specifically, private consumption is closely linked to short-term interest rates, highlighting the mechanism of budgetary shock transmission. An increase in inflation due to tax revenue shocks leads to an uptick in short-term interest rates, prompting economic agents in an uncertain environment to curtail their consumption.

A positive shock to tax revenues yields a significantly positive response in short-term interest rates. Increased tax rates result in reduced asset demand in an uncertain economy, leading to a decline in long-term interest rates and an increase in short-term interest rates.

A positive shock to government expenditures generates a significant and positive effect on short-term interest rates and borrowing rates linked to public debt. Higher expenditures lead to growth in current public debt and a reduction in its stock, contributing to an increase in interest rates. This dynamic aligns with Giovannini and De Melo's theory (1993). Furthermore, short-term interest rates react more strongly to tax revenue shocks than to government expenditure shocks (Lachaine 2017). This observation can be explained by the shock transmission mechanism. According to this economic logic, decisions regarding tax revenues stem from expenditure decisions. Therefore, an increase in tax revenues aimed at financing expenditures tends to mitigate the positive impact generated by expenditure shocks on short-term interest rates, resulting in a relatively subdued effect (Jørgensen, Ravn 2022).

The economic uncertainty index exhibits a significant negative reaction to tax revenue shocks, while a positive and significant reaction is observed following government expenditure shocks. This observation can be interpreted by considering the determination of the economic uncertainty index, which is largely influenced by the pace of fiscal measures (Favero, Giavazzi 2007: 4–7). Thus, an increase in tax revenues signifies more stable fiscal governance (by definition), contributing to reduced uncertainty associated with economic policy. Conversely, an increase in non-productive consumption expenditures leads to heightened economic uncertainty (OECD 2011).

A positive shock to tax revenues triggers significantly positive short-term production reactions. This response is explained by the revenue increase, subsequently leading to capital expenditure growth (Akpan 2005: 51–69). Conversely, a positive shock to consumption expenditures induces a decrease in production. This decreases stems from increased non-productive expenditures, such as defence-related expenses, resulting in increased tax pressure to cover their magnitude. This tax pressure leads to a decline in aggregate demand, translating into reduced production (OECD 2011).

The impulse response plots of variables to budgetary shocks exhibit substantial coherence with previous results, considering that the significance of the curves relies on trends coherent with economic theory’s expectations and prior knowledge (Hu, Sanyal 2016).

6. Conclusion

This article aims to identify the response of the inflation rate to fiscal shocks in the Russian economy during the period 2003T1-2022T4 while recognising the importance of the uncertaineconomic environment in which budgetary decisions are taken and their effects on inflation. To do this, we have focused on government spending, including military expenditure, and total revenue to estimate a two-part VAR(S) model. By introducing the economic uncertainty index in the (S)VAR model, we found that the estimate of fiscal shocks on inflation is statistically significant. Statistically, this indicates that variations in economic uncertainty influence the relationship between fiscal shocks and inflation. Economically, this implies that economic uncertainty can change the behaviour and political decisions of economic actors, thereby affecting the transmission of fiscal shocks to inflation.

However, when the economic uncertainty index is excluded from the VAR(S) model, estimates become statistically insignificant, which means that the effects of fiscal shocks are not properly captured. Statistically, economic uncertainty plays a crucial role in revealing the underlying economic dynamics and providing key clues on how to grasp these shocks. Economically, this suggests that increases in government spending to boost demand may be ineffective in a context of high uncertainty, as consumers may be reluctant to react immediately to budgetary measures due to a lack of clarity. Similarly, the impact of tax revenue increases on inflation could be mitigated by uncertainty, with economic operators reacting differently to tax adjustments depending on their perception of the economic future.

An interesting observation by the authors of this study is that the effects of fiscal shocks in Russia do not seem to significantly disrupt the level of inflation. One possible explanation lies in the robustness of the Russian economy, which is heavily dependent on revenues from oil and gas exports to Europe. During the first months of the Ukrainian conflict, these exports increased by about $47 billion, according to the country's finance ministry.

Some recommendations for Russian economic policies and future research may be expressed by:

- Integration of the Economic Uncertainty Index: to improve the effectiveness of budgetary policies, it is essential to integrate the economic uncertainty index into their formulation. This will help to better understand the reactions of economic actors and optimise the impact of government spending on aggregate demand, especially in times of high uncertainty.

- Strengthening communication and transparency in economic policies: improved communication and transparency of economic policies can reduce economic uncertainty. The Russian authorities should provide clear and regular information on the planned budgetary objectives and measures. This includes transparent economic forecasts, evaluation reports of budgetary policies, and public consultations. Better communication could help stabilise business and consumer expectations, thereby promoting a faster and more effective response to budgetary measures.
- Using high-frequency data and extending the analysis to other countries: To deepen this understanding, future research could further explore the complex interactions between economic uncertainty, fiscal shocks, and other macroeconomic variables through the use of high-frequency data applicable to other countries, which could enrich that understanding and generalise our idea.

In summary, the introduction of the Economic Uncertainty Index in our model has enabled us to capture its moderating effect on the relationship between fiscal shocks and inflation in Russia, improving our understanding of how the changes in economic uncertainties affect the transmission of shocks and, in particular, how they mitigate the rate of inflation. Each time, some recommendations are proposed.

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Appendices

Appendix (A)

Table (A.1). Optimal Lag order selection

Lag	AIC	SC	HQ
0	-9.658	-9.395	-9.569
1	-11.180	-9.496*	-10.471
2	-11.646	-8.300	-10.315
3	-14.058	-9.150	-12.106
4	-14.718	-8.248	-12.145*
5	-14.886	-6.856	-11.693
6	-15.079*	-5.486	-11.264

Source: Authors'compilation. (*) Indicates the selected optimal lag order by the information criteria.

Table (A.2). Residuals normality tests

Skewness	Kurtosis	Jarque-Bera
(0.1179) [11.5089]	(0.1124) [11.6584]	(0.0776) [23.1624]

Source:Authors' compilation. ( ) P-value and [ ] statistic. P-value > 0 leads us to reject the null hypothesis of non-normality in residuals.

Table (A.3).LM test for error autocorrelation

Lag	LRE* stat	df	Prob.	Rao F-stat	df	Prob.
1	54.83356	49	0.2630	1.140773	(49, 85.7)	0.2933
2	69.54737	49	0.0883	1.557650	(49, 85.7)	0.0964
3	54.10241	49	0.2859	1.121487	(49, 85.7)	0.3170
4	68.98799	49	0.3814	1.140752	(49, 85.7)	0.4800
5	64.59247	49	0.0669	1.410946	(49, 85.7)	0.0815
6	31.88112	49	0.9723	0.592939	(49, 85.7)	0.9759
7	70.22240	49	0.3250	1.578156	(49, 85.7)	0.3923

Source: Authors'compilation. Null hypothesis: The series are not correlated at the p level and are rejected if Prob <0.05.

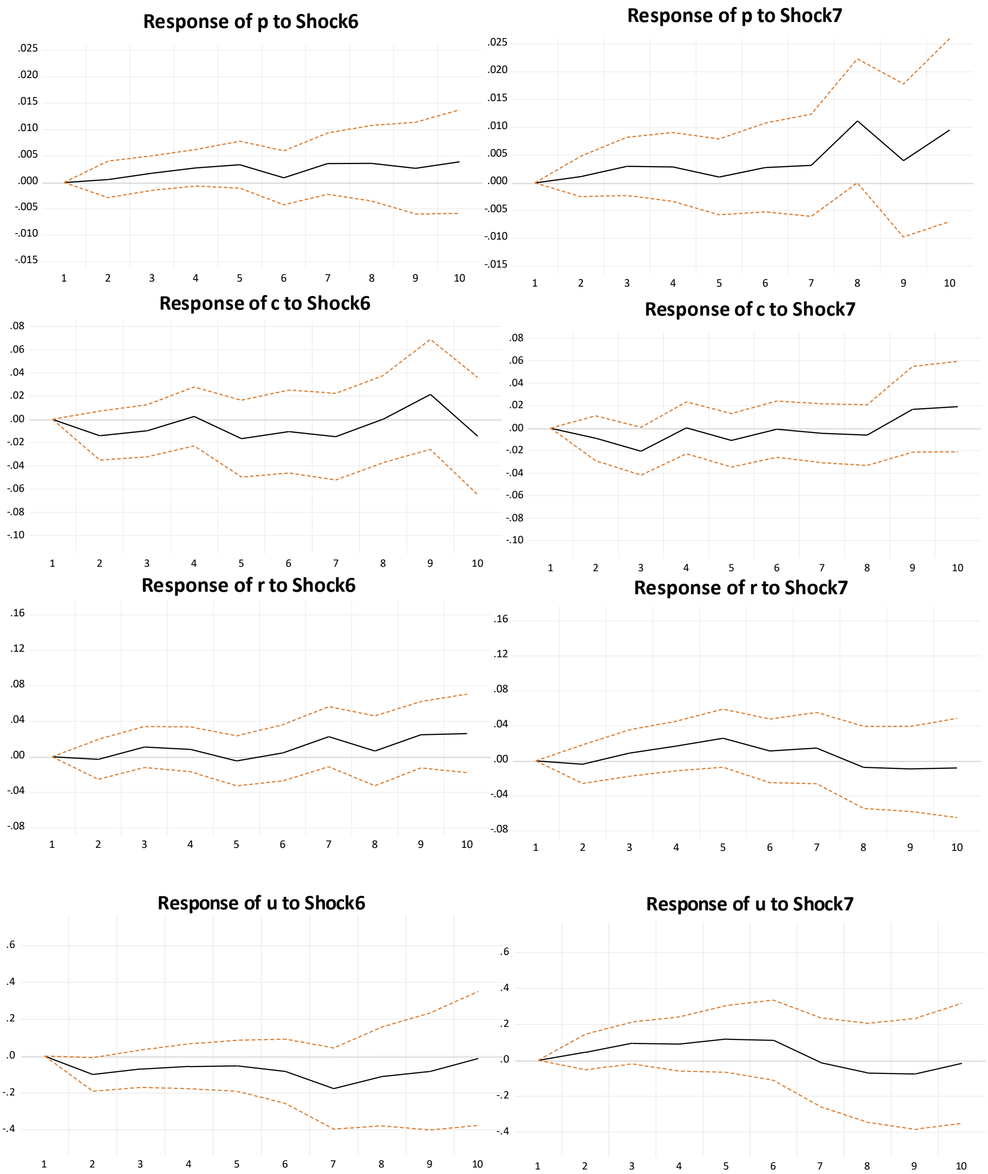
Appendix (B)

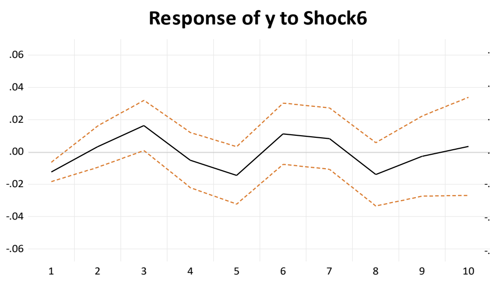
Table (B.1).(S)VAR model estimation

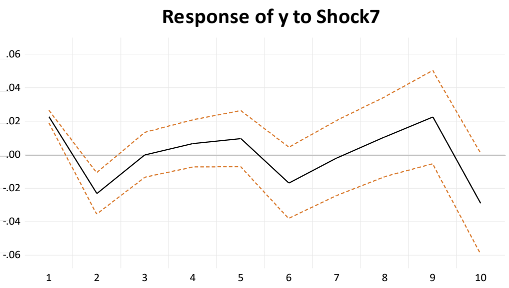
Coefficient	Std. Error	z-Statistic	Prob.	Coefficient	Std. Error	z-Statistic	Prob.
0.757711	0.654326	1.158002	0.2469	1.302698	0.737316	1.766812	0.0896
-2.426556	0.664199	-3.653358	0.0003	-2.777178	0.657863	-4.221513	0.0145
4.444862	2.082123	2.134774	0.0028	-0.005973	0.064782	-0.092201	0.8901
-0.541817	0.271510	-1.995567	0.0399	1.778604	0.194510	9.144027	0.0061
1.536655	0.585079	2.626407	0.0039	0.023784	0.052722	0.451123	0.4005
1.224578	0.245759	4.982840	0.0000	-3.781255	0.351045	-10.771422	0.0009
0.239743	0.118531	2.022609	0.0107	2.182755	0.300860	7.255044	0.0018
1.517995	0.507558	2.990783	0.0028	-0.000588	0.000682	-0.861682	0.2397
-0.359081	0.048018	-7.478039	0.0000	-0.007831	0,001203	-6.512113	0.0045
-0.772601	0.135368	-5.707428	0.0000	-0.004988	0,001400	-3.562701	0.0177
-0.019799	0.067301	-0.294186	0.7686	-1.008407	24,699513	-0.040827	0.8861
-1.739041	0.502294	-3.462200	0.0005	0.000663	0,000095	7.007823	0.0005
0.253003	0.048405	5.226816	0.0000	-0.002155	0,000229	-9.427312	0.0054
0.283251	0.050443	5.615256	0.0000	-0.007355	-0,000700	10.502594	0.0002
0.260755	0.120239	2.168644	0.0167	0.010702	0,000892	12.00000	0.0000
-0.026300	0.010515	-2.501096	0.0124	0.059418	0.004952	12.00000	0.0000
-2.777178	0.523186	-5.308204	0.0000	0.059761	0.004980	12.00000	0.0000
0.493752	0.219678	-2.247619	0.0247	0.254707	0.021226	12.00000	0.0000
0.787604	0.249253	3.159863	0.0016	0.022726	0.001894	12.00000	0.0000
-0.299024	0.109723	-2.725266	0.0218	0.048066	0.004006	12.00000	0.0000
0.152578	0.048618	3.138322	0.0017	0.019829	0.001652	12.00000	0.0000

Source: Authors'compilation. Note: The significance of autoregressive elasticity if prob <0.05.

Figure (B.1).Responses to budgetary structural shocks in the (S)VAR model







Source: Authors'compilation. Note: shock 6 = a fiscal revenue shock; shock 7 = a government expenditures shock.

Appendix (C)

Table (C.1).Optimal lag order selection for the model estimated, excluding the '*u*' variable

Lag	AIC	SC	HQ
0	-11.38773	-11.19501	-11.31118
1	-12.04241	-10.69331	-11.50653
2	-12.55955	-10.05408	-11.56435
3	-15.07394	-11.41210*	-13.61941*
4	-15.46561	-10.64741	-13.55176
5	-15.41095	-9.436378	-13.03778
6	-15.46579	-8.334852	-12.63330
7	-16.62894*	-8.341628	-12.33712
8	-14. 65739	-8.362854	-12.54128

Source: Authors compilation. (*)Indicates the selected optimal lag order by the information criteria.

Table (C.2). Residual normality tests for the model estimated, excluding the '*u*' variable

Skewness	Kurtosis	Jarque-Bera
(0.1326)	(0.5264)	(0.2441)
[9.815495]	[5.136531]	[14.95203]

Source: Authors'compilation. ( ) P-value and [ ] statistic. P-value > 0 leads us to reject the null hypothesis of non-normality in residuals.

Table (C.3).LM test for the model estimated, excluding the ‘*u*’ variable

Lag	LRE* stat	df	Prob.	Rao F-stat	df	Prob.
1	31.60282	42	0.6778	0.853075	(42, 73.6)	0.6954
2	41.37039	42	0.2477	1.183312	(42, 73.6)	0.2678
3	22.24488	42	0.9648	0.568584	(42, 73.6)	0.9677
4	45.76537	42	0.1276	1.344001	(42, 73.6)	0.1423
5	38.25361	42	0.3675	1.074017	(42, 73.6)	0.3896
6	42.91570	42	0.1989	1.238919	(42, 73.6)	0.2173
7	35.88579	42	0.4740	0.993480	(42, 73.6)	0.4960
8	23.64619	42	0.9436	0.609326	(42, 73.6)	0.9480

Source: authors’compilation. Note: Null hypothesis: The series are not correlated at the p level and are rejected if Prob <0.05.

Table (C.4).(S)VARmodel estimation, excluding the ‘*u*’ variable

Coefficient	Std Error	z-Statistic	Prob	Coefficient	Std Error	z-Statistic	Prob
0,005021	0,002838	1,769832	0,0623	-0,393231	0,045894	-8,568331	0,0092
-0,456741	1,305552	-0,349826	0,3256	0,008538	0,004373	1,952623	0,0532
-2,915562	0,408877	-7,128012	0,0002	-2,362894	1,243856	-1,899652	0,1294
-1,570430	0,895424	-1,753968	0,0789	1,002158	0,309330	3,239768	0,0068
0,086582	0,165595	0,522633	0,1173	-0,154647	0,125222	-1,234986	0,2630
-0,729856	0,236638	-3,085514	0,0014	-0,007821	0,000804	-9,721627	0,0013
-2,303921	1,263115	-1,823999	0,1888	-1,004393	0,518378	-1,937568	0,0992
-0,880346	0,538320	-1,635359	0,4790	3,865082	14,666963	0,263523	0,7526
-5,991072	7,071389	-0,847227	0,7903	0,077788	0,036872	2,109652	0,3754
-1,004393	0,507893	-1,977568	0,0992	2,210508	0,464159	4,762398	0,0073
3,082865	8,742666	0,352623	0,7932	-0,641457	0,323585	-1,982346	0,3056
2,770788	1,378740	2,009652	0,4794	-2,820071	-3,907934	0,721627	0,0047
-1,407686	0,485471	-2,899632	0,0231	-5,390043	2,781860	-1,937568	0,0877
0,015889	0,005987	2,653859	0,0156	2,088652	7,925881	0,263523	0,2657
-0,094837	0,051374	-1,846017	0,0933	0,057892	0,030300	1,910652	0,5436
0,108158	0,062211	1,738566	0,1263	0,026701	0,115183	11,00000	0,0000
0,097918	0,277685	0,352623	0,2329	1,770925	0,089087	11,00000	0,0000
3,667023	1,886666	1,943652	0,0562	0,050001	0,160993	11,00000	0,0000
4,035792	15,745193	0,256319	0,5216	0,049375	0,136364	11,00000	0,0000
-0,084662	0,081724	-1,035947	0,3791	0,079386	0,162988	11,00000	0,0000
0,097211	0,077403	1,255907	0,3203	0,070723	0,064294	11,00000	0,0000

Source: Authors’compilation. Note: The significance of autoregressive elasticity if prob<0.05.

Figure (C.1).Responses to budgetary structural shocks in the SVAR model, excluding the variable '*u*'

Source: Authors'compilation. Note: shock 6 = a fiscal revenue shock; shock 7 =a government expenditures shock.

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Exchange rate behaviour in ASEAN countries – a sensitivity analysis

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Aim: The study examined the behavior of exchange rate in ASEAN countries. This was highly necessitated in order to account for the structural break in the data set occasioned by global financial crisis.

Research method: The quantile regression sensitivity analysis was performed on daily series of exchange rate volatility for 8 ASEAN countries having divided our sample into two, before and after the financial crisis eras. Periods of low market volatility (2001–2006 plus 2010–2017) and high market volatility (1990–2000, 2007–2009, plus 2018–2023) correlate to the periods before and after the financial crisis, respectively.

Findings: The empirical finding going forward is that since the global financial crisis took effect, exchange rate volatility has not been effectively curtailed by the governments and monetary authorizes of ASEAN countries especially in Thailand, Malaysia, Indonesia and Vietnam respectively. There is therefore the need for a policy fight in favour of stability of the currency exchange rates.

Originality: The originality of the research resides with the sensitivity analysis which validates the presence of high persistence in the volatility of the Thai Baht exchange rate throughout the quantiles. This was followed on by the high persistence in the exchange rate of the Malaysian ringgit which began at the 70th quantile in the pre-financial crisis period with a persistence value of 1.0097 as against the 30th quantile in the post-financial crisis estimations with a persistence value of 1.0387. The Indonesian Rupiah and Vietnamese dong took turns as regards volatility persistence. We also found significant ARCH effect which instigated further estimations of the GARCH and FIGARCH models as robustness checks.

Contributions: With the GARCH results, the study contributed to establishing persistence of volatility in the exchange rates of all ASEAN countries in our sample, with varying degrees and this could be attributed instabilities in the economies. Explicitly, the significance of the FIGARCH coefficient confirms the persistence of volatility over time with considerable long-term memory effect. This implies that once the exchange rate becomes volatile, such volatility last long, influencing future volatility levels noticeably in all the countries. Exchange rate volatility persistence of the Singapore Dollar was very low.

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JEL: A20, B34, C50

1. Introduction

Exchange rate, being the value of one currency for the conversion to another, is influenced by numerous factors such as inflation, interest rates, oil price variation, growth of money in circulation, and income growth rate etc. (Umoru, Abugewa-Ejegi, Effiong 2023; Umoru, Akpoviroro, Effiong 2023). The study aimed at evaluating the behaviour of exchange rate in ASEAN countries. The members of the Association of Southeast Asian Nations (ASEAN) covered in this research include Indonesia, Singapore, Philippines, Thailand, Vietnam, Cambodia, Myanmar, and Malaysia. Price and exchange rate stability is the core objective of the monetary authorities of these countries. Indonesia central bank operates a free floating exchange rate regime. Hence, Indonesia rupiah exchange rate is strongly swayed by capital flows and exports earnings. The monetary policy framework of Singapore is exchange rate-centered. The Singapore dollar is regulated against a basket of currencies whose composition is reviewed occasionally in order to accommodate changes in trade patterns. The trade-weighted exchange rate fluctuates within a policy band that is fixed nearby a targeted level and a given appreciation rate. The policy adjustments to the parameters of the exchange rate band are announced every three months. The Philippine government implements the floating exchange rate system for the Philippine Peso. Accordingly, anytime foreign shocks disturb the domestic economy, the flexible peso exchange rate serves as an automatic stabilizer that regulates and provides a restoration to macroeconomic balance. The Bank of Thailand operates a managed float exchange rate system whereby market mechanism fixes the value of the Thai baht while the Bank of Thailand intervenes when the Thai baht exchange rate is extremely volatile.

The government of Vietnam operates two exchange rate policies, namely, “following” the market and “unifying” the market. In Cambodia, volatility in the exchange rate, the National Bank of Cambodia (NBC) committee meets to strategize to bring the exchange rate back within range. Currently, the Central Bank of Myanmar

(CBM) issues daily official parallel FX rates having aggregated transactions reported by commercial banks to the online forex trading platform. The exchange rates are accessible and the CBM upholds control over the rates used in these transactions. In Malaysia, the Bank Negara Malaysia (BNM) operates a floating exchange rate system since 2016. The BNM permitted exporters to exchange 75% of their proceeds into the ringgit, while 25% of the proceeds are retained in foreign currency. This practice relaxed forex hedging restrictions and created some liberalization measures which has amplified the volatility of the Malaysian ringgit because huge flexibility is allowed for exporters. The behavior of the exchange rates of ASEAN currencies is of policy significance because financial traders and marketers can rightly predict the performance of the foreign exchange market in ASEAN countries and factor in the volatility of the exchange rate when making investment decisions. As a result, these investors are guided in their decision-making. Overall, the research findings are relevant in that they provide policy guidelines to asset portfolio decision-makers and forex traders in the ASEAN financial markets. In summary, the stability of the exchange rates of all ASEAN currencies relative to foreign currencies is essential in order to achieve and maintain price stability and stability in the financial market. The study is of immense value to policymakers who stand to benefit profoundly from policy findings as regards the formulation of effective regulatory policies; ideas into how digital currencies interact with traditional economic indicators can guide the development of regulatory frameworks that foster innovation while mitigating potential risks. Policymakers can use the findings to design measures that promote financial stability and ensure the responsible integration of forex markets within the broader economic landscape of all ASEAN nations. The next section reviews related and relevant literature. Section three discusses the research methodology, while section four discusses the results and policy implications. The study concludes with section five.

2. Literature review

Kipkorir & Mutai (2024) focused on the influence of commodity price fluctuations on the behavior of the Kenyan Shilling against major currencies from 2015 to 2023. The study sourced its exchange rate and commodity price data from the Central Bank of Kenya and global commodity exchanges. Employing VECM, the research aimed to uncover how global changes in commodity prices, especially tea and coffee, impact the exchange rate. The findings indicated a significant and lasting impact of commodity price volatility on the exchange rate, with a significance level less than 0.05. Long-term analysis showed a slow adjustment to equilibrium, with an annual speed of 3.7%, reflecting the protracted effect of commodity price changes on the currency. The study suggest that Kenya's central bank should enhance its monitoring of commodity markets and potentially engage in futures contracts to hedge against predictable fluctuations in currency values. Okonkwo & Mbekeani (2023) investigated the behavioral patterns of the Rand against the US Dollar during periods of political instability from 2010 to 2022. Using exchange rate data from the South African Reserve Bank, the study utilized a combination of unit root tests to ensure data stationarity and co-integration analysis to establish relationships, followed by a VECM to estimate the dynamics. The research found that political instability leads to significant volatility in the Rand, with exchange rate movements showing heightened sensitivity during election cycles and major political announcements. The significance level for these fluctuations was found to be below 0.05, indicating a strong relationship. The study also revealed that this relationship persists in the long run, with a speed of adjustment to equilibrium of 5.4% annually. In the short run, each 1% increase in political instability could lead to approximately a 0.12% increase in exchange rate volatility. The study recommended that policymakers and investors consider the timing of political events when assessing currency risk, suggesting that strategic currency management could mitigate the adverse effects on the Rand.

Djouakaa et al. (2023) examined real effective exchange rate causes. Their results demonstrated that money supply; direct investment, inflation, imports and interest rate are the main determinants of the real effective exchange rate in the franc zone after using the Driscol-Kraay method on panel data. Also, the use of the Panel Corrected

Standard Error method also revealed the same results. In addition, the analysis of the specificities of the 15 Franc Zone countries made it achievable to take into account the heterogeneity of the panel. As a result, the central banks of each of the 15 Franc Zone countries must incorporate the consequences of the rate of exchange when formulating their monetary policies. Moreover, each government, when formulating its macroeconomic policy, must take into account the repercussions of the exchange market. Aizenman et al. (2023) investigated the link between RER behaviour and international reserve in the era of financial integration. They utilized nonlinear regressions and panel threshold regressions techniques to determine whether international reserve is a determinant of real exchange rate. Their study covered over 110 countries making use of panel data from 2001-2020. Some of the countries include Algeria, Botswana, Brazil, Guinea, Honduras, Hungary, Guinea, Honduras, Hungary and others. The findings show that term of trade shocks had significant on real exchange rate. The results also indicate that countries with intermediate levels of financial development will have a more powerful buffer.

The study conducted in SSA nations with a focus on the oil-importing and exporting nations by Korley & Giouvris (2022) evaluated the influence of oil price and oil volatility index on the exchange rate. Their study employed quantile regression and Markov switching models to evaluate their joint effect and showed that oil volatility considerably influenced the exchange rate of all countries. They established that the rising and falling oil prices leads the local currency to devalue or appreciate. They submitted that exchange rates respond to oil price and oil volatility mostly at lower quantiles for all countries which indicates the sensitivity of investors to risks and returns. Bangura et al. (2021) aimed to investigate the behavior of the Leone/US dollar exchange rate based on the influence of global oil price shocks in Sierra Leone throughout the post-war period from June 2002 to May 2020. Among the three estimated models, the EGARCH (1, 1) model emerged as the most suitable fit, with all mean and variance coefficients deemed significant. The empirical findings indicated that a rise in oil prices corresponded to a depreciation of the exchange rate in Sierra Leone amongst others.

Raksong & Sombatthira (2021) reported a positive impact on the REER of the ASEAN countries by some variables which include the ratio of FDI to GDP and government spending. Trade opening also had a substantial positive impact on REER in all the ASEAN countries studied except Vietnam while terms of trade had significant impact on REER in Malaysia, Philippines and Indonesia. Foreign direct investment also had significant impact on REER, but only in Vietnam. International reserve was shown to have long-run impact on REER in Malaysia, Thailand and Vietnam. Ani & Mashood (2021) undertook a comprehensive analysis to evaluate the behaviour of real exchange rate (RER) in Nigeria over a sample size of 60 years from the period of 1960-2020. The study utilized a multivariate co-integration test, he ADF and KPSS stationarity test as well as the VECM to analyze the data set. The result of the stationarity test showed that the macroeconomic variable under study had no stochastic trends and were stationery at all levels while the result of the Granger causality test showed that real GDP growth rate, inflation rate, money supply growth rate and government expenditure exerted significant influence on real exchange rate behavior.

Damayanthi & Gunawardhana (2021) analyzed the behavior of the real effective exchange rate in Sri Lanka with a direct linkage to external sector stability, sampling data from the period from 2010 to 2019. The research employed the vector auto regression (VAR) model as an analytical tool. The results of the study revealed that exchange rate behaved as theoretically expected with changes in policy rates of United States. Although, behavior of inflation, domestic interest rates and net exports were contrary to theoretical expectations. The findings show that even though domestic currency depreciation may lead to increased net exports, Sri Lanka, being a net importing economy, had suffered further depreciation of its domestic currency. Kalaj & Golemi (2020) focused their study on the bases for real exchange rate behaviour in Albania; sing the VAR model to assess data for the period of 1995-2015. The study aims to find the long run relationship among variables. The findings suggest that real exchange rate behavior can be decreased by increasing fiscal policies and decreasing monetary policies. Findings from the study also indicated a long-run relationship between real exchange rate behavior and trade.

Romo & Gallardo (2020) explored the bases for the behavior of RER behavior, particularly analyzing the effect of share of wages in output on RER. The study modeled the behavior RER of the domestic currencies of three countries: Mexico, Korea and France against the US dollar. The econometric technique, VAR model, was specified to find the long-run associations of the RER for each country. The results show that, for each country, there was a negative relationship between the RER and wage share. Findings also showed that Real Exchange Rate is positively related with labor productivity in France and Korea, but inversely related in Mexico. An explanation as to why RER tended to return to the long run normal value was also given. To Kahsay & Patena (2020), estimates from vector error correction models (VECMs) and vector autoregression models (VARs) indicate that the REER responds minimally to changes in fundamentals, with small and delayed responses. Three factors are proposed to have contributed to this minimal response: labor market conditions, price management, and remittance outflows. The policy implications suggest that labor market reforms, price liberalization, and policies encouraging domestic investment could improve the REER's response to economic fundamentals. Hien et al. (2020) undertook a comprehensive analysis on how huge amounts of remittance can have an effect on a country's Real Exchange Rate (RER) behavior, thereby causing the Dutch disease. The study focused on countries which receive a quite high value of remittances, specifically on 32 Asian developing countries. Using data covering the period from 2006 to 2016, the S-GMM for the linear dynamic panel data (DPD) was used to analyze the relationship between REER and remittances. The findings show that as per capita remittances increases by 1 percent, REER increases by 0.103 percent, signifying the existence of the Dutch disease. The results further indicated that in countries with low remittance to Gross Domestic Product, remittances results in REER appreciation while for countries having a ratio higher than 1 percent, higher remittances results in real effective exchange rate appreciation. Moreover, the study corroborates other findings why postulate that a flexible exchange rate regime leads to the dampening of the appreciation of the REER caused by increased remittances.

Hassan et al. (2020) did a study for 22 OECD countries. The fixed effect model was the estimated technique used to analyze the panel data covering the period 1980-2015. Their research amongst other results, reveals that countries with a large share of a working population tend to have an appreciating effect on the RER because increase in the working age population leads to an increase in the marginal product of capital which increases foreign direct investment thereby causing an increase in capital inflow. The terms of trade also had a significant effect on RER. The implications of the findings suggest that for developing countries experiencing a rapid aging population, there will be a negative effect on their international competitiveness due to RER appreciation. To counter this, the government of these countries will have to adopt a policy of saving more for the future by increasing the effective retirement age or through an increase in budget surplus.

After reviewing the literature, a gap is found: many scholars have not felt the need to conduct an empirical investigation into the behavior of ASEAN currency exchange rates within the context of a sensitivity analysis that considers financial crisis periods. Sensitivity analysis receives empirical focus in this study. Thus, trustworthy conclusions that direct the decision-making process regarding firms, investments, and the entire economy are reached by assessing the sensitivity of our data on the currency exchange rates of ASEEAN nations with respect to times of high and low volatility. Sensitivity analysis of this kind offers factual proof of the validity of study conclusions about the behavior of exchange rates in ASEAN nations. The current study focused on the ASEAN countries of Malaysia, Indonesia, Singapore, the Philippines, Thailand, Vietnam, Cambodia, Myanmar, and Vietnam.

3. Methodology

To analyze the behavior of exchange rates in ASEAN countries, we estimated quantile regression analysis on the daily series of exchange rate volatility for 8 ASEAN countries using daily data. This was highly necessitated in order to control for the structural break in the data set caused by global financial crisis. Within the scope of this research, 1990 to 2023, the following phases of structural breaks are discernible. The period from 1990 to 1991 witnessed low market volatility due to political stability that existed across countries, fiscal discipline on the part of the governments, absence of terrorism, etc. The period of 1990 to 2000 was a period of high market volatility due to Harshad Mehta Scam of 1992 in Indian that crashed the stock market and Asian tiger Financial crisis of 1997-1998 which led to the collapse of the Thai baht, the baht was floated resulting in a huge devaluation that spread to much of East Asia, also hitting Japan, as well as a huge rise in debt-to-GDP ratios. The period, 2001 to 2006 was characterized by low market volatility. This was occasioned by the fairly stable global exchange rate regime, global peace, and unified exchange rate system. The period, 2007 to 2009 saw high market volatility due to global financial crisis which expanded into a global banking crisis following the catastrophe collapse of investment bank Lehman Brothers in September 2008, stock market crashes, credit crunches, the bursting of financial bubbles, sovereign defaults, and global liquidity/currency crises etc. The period, 2010 to 2017 was branded by low market volatility as a result of global political stability and stability in prices of oil but individual economics had peculiar issues to deal with. The period, 2018 to 2023 was a period of high market volatility due to Russian-Ukraine war, US-China trade war, unrest from middle east, sustained global terrorism and outbreak of diseases, namely, the covid-19 pandemic and the fact that each country had to tackle other endogeneous issues. Hence, we divided our sample into two, namely, the pre-financial crisis sample and the post-financial crisis sample. This yielded a sensitivity analysis that provided evidence of the robustness of our regression estimates for exchange rate volatility persistence in ASEAN countries, namely, Indonesia, Singapore, Philippines, Thailand, Vietnam, Cambodia, Myanmar, and Malaysia. Accordingly, by deploying the quantile regression methodology to conduct the sensitivity analysis of the persistence in the volatility of exchange rate for the different countries in our sample across different quantiles, we contributed to empirical knowledge on exchange rate volatility persistence. Following the works of Benoit & Van Der Poel (2009), we specify the quantile autoregressive equation for exchange rate persistence by the p-th order AR process with random coefficient as in equation (1):

$$NEXC_t = \beta_0(\phi_t) + \beta_1(\phi_t)NEXC_{t-1} + \ldots. + \beta_p(\phi_t)NEXC_{t-p} \qquad (1)$$

Where β_i's are unknown parameters and (ϕ_t) is a sequence of independently distributed uniform random variables. From regression (1), the nominal exchange rate persistence (ρ) was measured as the sum of autoregressive coefficients (SARC). The standard SARC exchange rate persistence model is specified as follows:

$$\Delta NEXC_t = \varphi + \sum_{i=1}^{q-1} \beta_i NEXC_{t-i} + (\rho - 1) NEXC_{t-1} + v_t \qquad (2)$$

By equation (2), changes in the nominal exchange rate can be calculated as in equation (3):

$$\Delta NEXC_t = NEXC_t - NEXC_{t-1} \tag{3}$$

Making provision for intercept and trend terms in the SARC equation (2), we have as follows:

$$NEXC_t = \varphi + \phi_t + \rho NEXC_{t-1} + \sum_{i=1}^{q-1} \beta_i \Delta NEXC_{t-i} + (\rho - 1) NEXC_{t-1} + v_t \tag{4}$$

Accordingly, based on the SARC methodology, the exchange rate persistence could further be represented mathematically as:

$$\rho = \beta_1 + \beta_2 + \ldots\ldots. + \beta_p = \textstyle\sum_i^n \beta_p \tag{5}$$

Expressing the τth conditional quantile function as a function of $NEXC_t$ from equation (5), we have:

$$Q_{NEXCt}\,(\tau/S_{t-1}) = \beta_0(\tau) + \beta_1(\tau)NEXC_{t-1} + \ldots\ \beta(\tau)NEXC_{t-p} \tag{6}$$

Since $Q_{NEXCt}\,(\tau/S_{t-1})$ is the conditional distribution function, equation (6) becomes the baseline quantile autoregressive (p) model of exchange rate persistence. Our vector of predictor variables is constructed as:

$$Z_t = (1, NEXC_{t-1}, NEXC_{t-2} \ldots, NEXC_{t-p}), \tag{7}$$

Hence, the quantile autoregressive (p) model of exchange rate persistence can be specified as:

$$NEXC_t = Q_{NEXCt}(\tau/S_{t-1}) + e_t = Z_t'\beta_\tau + e_t \tag{8}$$

Where β_τ is the autoregressive quantiles and its estimate is given by:

$$\beta_\tau^* = \arg_{\beta \in R} \min_\psi \sum_{t=1}^{T} \rho_\tau (NEXC_t - Z_t'\beta) \tag{9}$$

The quantile loss function is defined as:

$$\rho_\tau(NEXC_t) = NEXC_t[\tau - I(v < 0)] \tag{10}$$

The linear function of parameters is characterized in equation (9). For the investigation of persistence in the exchange rate, volatility in the exchange rate is adjudged persistent if $\rho \geq 1$. In effect, a higher sum of AR coefficients signifies high volatility persistence. In addition to the quantile regression estimations, we further tested for the ARCH effect and thereafter embarked on the estimation of the GARCH/FIGARCH model estimations. The ARCH model based on the maximum likelihood estimation approach, specifically an ML ARCH model with a normal distribution, which optimizes the coefficients using BFGS/Marquardt steps. The significance of the ARCH effect guaranteed further estimations using the FIGARCH model. Estimating financial models using both quantile and FIGARCH-DCC methods involves navigating the complexities of conditional quantile and dynamic correlations. In quantile Regression, various quantile of the response variable is modeled, offering flexibility in handling non-normally distributed data. The estimation process involves minimizing weighted absolute deviations through optimization algorithms. FIGARCH-DCC, combining Fractionally Integrated GARCH for volatility and Dynamic Conditional Correlation for inter-asset correlations, requires careful parameter estimation. The models can be integrated through joint or sequential estimation, with considerations for computational intensity and model diagnostics. In the joint estimation approach, parameters of both quantile regression and FIGARCH-DCC are estimated simultaneously, demanding specialized software and computational resources. Alternatively, in sequential estimation, FIGARCH-DCC parameters are estimated. The daily exchange rate data that spanned the period from January 1, 1990 to December 30, 2023 plus weekends and holidays were utilized in this research. The before and after the financial crisis periods correspond to periods of low market volatility (2001–2006 plus 2010–2017) and high market volatility (1990–2000, 2007–2009, plus 2018–2023), respectively. Data on the 8 ASEAN countries, namely, Indonesia, Singapore, Philippines, Thailand, Vietnam, Cambodia, Myanmar, and Malaysia were sourced from databases of the World Trade Organization and World Bank.

4. Results

This section deals with the presentation of results, analysis and interpretation of the results. As a robustness check, we further estimated persistence in the volatility of exchange rate across different quantiles for the currency exchange rates of all the countries using the quantile regression algorithm. Table 1 reports the quantile results for exchange rate volatility persistence before the global financial crisis while Table 2 provided the quantile results for persistence in the volatility of exchange rate. The most striking result was that obtained for Thailand where exchange rate volatility persistence exceeded unity for both samples. This implies high persistence in the volatility of Thai Baht exchange rate. This was followed by the high persistence in the exchange rate of the Malaysian ringgit which began at the 70^{th} quantile in the pre-financial crisis period with a value of 1.0097 as against the 30^{th} quantile in the post-financial crisis estimation with a persistence value of 1.0387. In Indonesia, exchange are volatility persistence emerges in the 80^{th} quantile with a persistence value of 1.0163 for the pre-financial crisis era while it dropped to the 30^{th} quantile with a persistence value of 1.0234 after the financial crisis. In Vietnam, persistence in the volatility of the exchange rate of the Vietnamese dong occurred at the 90^{th} quantile with a coefficient of 1.0254 before the financial crisis while such persistence took effect at the 70^{th} quantile with a value of 1.0921 after the financial crisis. Exchange rate volatility persistence was extremely low in Singapore for both periods of analysis. Thus, only the Singapore Dollar exchange rate was less volatile while the Thai Baht was the most volatile, followed by the Malaysian ringgit, Indonesian Rupiah, and Vietnamese dong. The implication of the sensitivity analysis is that since the global financial crisis took effect, exchange rate volatility has not been effectively curtailed by the governments and monetary authorizes of MENA countries especially in Thailand, Malaysia, Indonesia and Vietnam respectively.

Table 1. Quantile results for exchange rate volatility (before financial crisis)

Quantile	Indonesia	Singapore	Philippines	Thailand	Vietnam	Cambodia	Myanmar	Malaysia
Q1	0.3371	0.0594	0.2489	1.0019	0.4364	0.136	0.1386	-0.2734
Q2	0.3569	0.0601	0.2571	1.0457	0.4821	0.2573	0.1579	0.3258
Q3	0.4562	0.0696	0.27469	1.2865	0.5279	0.2846	0.2928	0.4920
Q4	0.5103	0.0718	0.38572	1.3091	0.6128	0.3509	0.3028	0.5103
Q5	0.6720	0.0790	0.44286	1.4270	0.6245	0.3610	0.4287	0.6231
Q6	0.6891	0.0712	0.50286	1.5622	0.7256	0.4928	0.5211	0.6357
Q7	0.8824	0.0832	0.6419	1.6890	0.8326	0.6192	0.6972	1.0097
Q8	1.0163	0.1795	0.8735	1.7028	0.9130	0.7355	0.8153	1.2236
Q9	1.3420	0.4026	1.0289	1.7311	1.0254	0.7392	0.8967	1.5810

Source: Authors' estimation results with Eviews 10.

Table 2. Quantile results for exchange rate volatility (after financial crisis)

Quantile	Indonesia	Singapore	Philippines	Thailand	Vietnam	Cambodia	Myanmar	Malaysia
Q1	0.5873	0.4280	0.3702	1.1389	0.55780	0.2793	0.2489	0.3485
Q2	0.6680	0.4455	0.4986	1.1472	0.5692	0.3459	0.3027	0.3690
Q3	1.0234	0.4557	0.50287	1.2350	0.6043	0.4428	0.4529	1.0387
Q4	1.0011	0.5139	0.6311	1.3347	0.6274	0.4703	0.5126	1.1374
Q5	1.2237	0.6630	1.0326	1.4655	0.7139	0.5920	0.5510	1.1520
Q6	1.2347	0.7241	1.0437	1.5609	0.7725	0.6397	0.6013	1.3292
Q7	1.4528	0.8251	1.2247	1.6234	1.0921	0.7355	0.7586	1.3379
Q8	1.5920	0.9428	1.3529	1.6235	1.5130	0.8230	0.9326	1.4261
Q9	1.6134	0.9631	1.4750	1.9352	1.6093	0.8250	0.9910	1.7250

Source: Authors' estimation results with Eviews 10.

In Table 3, the lagged value of the nominal exchange rate (*NEXC*) had a coefficient of 0.982322. This coefficient is highly significant with a z-statistic of 431.7296, suggesting that the current value of *NEXC* is heavily influenced by its immediate past value, with nearly a one-to-one relationship. The intercept (C) in the mean equation is 0.743140, also significant, indicating a consistent additive effect on the exchange rate movement per period. In the variance equation, which models the volatility (conditional heteroskedasticity) of the *NEXC*, we observe that the constant term (C) is 2.423849. This value, significantly different from zero (with a z-statistic of 53.68882), suggests a base level of variance in the *NEXC* that is quite high, reflecting inherent market volatility or economic instability. The coefficient for

RESID(-1)^2 is -0.017478, which is significant and negative (z-statistic of -42.27400). This indicates a mean-reverting volatility pattern; higher volatility in one period tends to be followed by lower volatility, a common characteristic in financial time series reflecting periods of market corrections following shocks. The significant persistence of real exchange rate and its volatility implications suggest that past values and volatility shocks have a strong predictive power on future values and volatility levels in Indonesia, respectively.

Table 3. ARCH results for Indonesia

Mean Equation				
Variable	Coefficient	Std. Error	z-Statistic	Prob.
C	0.743140	0.128465	5.784748	0.0000
NEXC (-1)	0.982322	0.002275	431.7296	0.0000
Variance Equation				
C	2.423849	0.045146	53.68882	0.0000
resid(-1)^2	-0.017478	0.000413	-42.27400	0.0000
F=190.357(0.000)				

Source: Authors' estimation results with Eviews 10.

Table 4. ARCH results for Singapore

Mean Equation				
Variable	Coefficient	Std. Error	z-Statistic	Prob.
C	0.373365	0.209037	1.786120	0.0741
NEXC(-1)	0.967368	0.014555	66.46428	0.0000
Variance Equation				
C	0.608470	0.023843	25.52022	0.0000
RESID(-1)^2	-0.015426	0.006147	-2.509593	0.0121
F=572.91(0.000)				

Source: Authors' estimation results with Eviews 10.

In Table 4, the coefficient for nominal exchange rate is 0.967368, which is extremely high and significant with a z-statistic of 66.46428, pointing to a strong persistence in the exchange rate. Essentially, this suggests that the current month's exchange rate is almost a direct continuation of the previous month's rate, which indicates a stable or slowly evolving exchange rate over time. The ARCH component specifies how volatility (variance) of the real exchange rate is affected by shocks (RESID(-1)^2) from the previous period. The constant is 0.608470, indicating a base level of volatility that is significant given a z-statistic of 25.52022. The coefficient for RESID(-1)^2 is -0.015426, which is statistically significant (z=-2.509593, p=0.0121), indicating a negative relationship between past squared residuals and current volatility. This negative value suggests a mean-reverting volatility behavior of exchange rate in Singapore; periods of high volatility tend to be followed by lower volatility; a typical characteristic observed in financial markets known as the volatility clustering.

Table 5. ARCH results for Philippines

Mean Equation				
Variable	Coefficient	Std. Error	z-Statistic	Prob.
C	0.171417	0.187246	0.915467	0.3599
NEXC(-1)	0.979605	0.017294	56.64583	0.0000
Variance Equation				
C	0.083953	0.001614	52.01721	0.0000
RESID(-1)^2	-0.005885	0.000477	-12.32860	0.0000
F=1347.2 (0.0000)				

Source: Authors' estimation results with Eviews 10.

Table 5 shows that the coefficient for real effective exchange rate is 0.979605, which is significant at the 0.05 level. Hence, real effective exchange rate in the previous period is a strong predictor of the real effective exchange rate in the current period, with a 1 unit increase in real effective exchange rate associated with approximately a 0.98 percent increase in the current *NEXC*, holding all other variables constant. The ARCH model results suggest that the real exchange rate in Philippines exhibits time-varying volatility that can be partially captured by its own past values. This could have implications for the demand for in Philippines.

From Table 6, the coefficient for nominal exchange rate lagged one-period is extremely significant (z-statistic of 59.96873) and near unity (0.976800), pointing towards a strong autoregressive behavior where current real effective exchange rate values closely follow past values. The variance equation, which models the volatility of the real effective exchange rate, features a small but highly significant constant (C=0.030006; z-statistic of 51.68566), indicating a base level of volatility that is consistent yet relatively low, given the scale of the coefficient. This suggests that external shocks or inherent economic volatility in Thailand's market is modest but persistent over time. The ARCH model's findings shows that real effective exchange rate in Thailand exhibits predictable behavior based on its historical values, its volatility is not overly influenced by its immediate past, except in the form of mean-reversion in response to shocks. This could imply that the Thailand exchange rate is relatively stable, with inherent mechanisms that dampen the impact of large fluctuations over time, potentially reflecting effective monetary policy interventions or a stable macroeconomic environment.

Table 6. ARCH results for Thailand

Mean Equation				
Variable	Coefficient	Std. Error	z-Statistic	Prob.
C	0.121671	0.101827	1.194875	0.2321
NEXC(-1)	0.976800	0.016288	59.96873	0.0000
Variance Equation				
C	0.030006	0.000581	51.68566	0.0000
RESD(-1)^2	-0.006750	0.000206	-32.74471	0.0000
F=486.17(0.0000)				

Source: Authors' estimation results with Eviews 10.

Table 7. ARCH results for Vietnam

Mean Equation				
Variable	Coefficient	Std. Error	z-Statistic	Prob.
C	0.373365	0.209037	1.786120	0.0741
NEXC(-1)	0.967368	0.014555	66.46428	0.0000
Variance Equation				
C	0.608470	0.023843	25.52022	0.0000
RESID(-1)^2	-0.015426	0.006147	-2.509593	0.0121
F=3002.3(0.0000)				

Source: Authors' estimation results with Eviews 10.

The ARCH model results for Vietnam on Table 7 shows 0.967368 coefficient of the real effective exchange rate Lagged one period. The coefficient is highly significant (z-statistic of 66.46428) and very close to one. This indicates that the current *nexc* value is almost perfectly predictive by its previous value, reflecting strong persistence or inertia in the exchange rate. This finding implies that once the exchange rate reaches a certain level, it is likely to remain near that level in the subsequent period unless significant economic events occur. The variance equation is critical for understanding the volatility of real effective exchange rate. The constant term (C) is 0.608470, highly significant (z-statistic of 25.52022), indicating a substantive inherent volatility in the *nexc*. This reflects ongoing economic volatility in Vietnam potentially due to fluctuating commodity prices, political uncertainty, or external economic shocks.

Table 8. ARCH results for Cambodia

Mean Equation				
Variable	Coefficient	Std. Error	z-Statistic	Prob.
C	0.095810	0.152964	0.626354	0.5311
NEXC(-1)	0.995107	0.008166	121.8662	0.0000
Variance Equation				
C	0.215612	0.005022	42.92938	0.0000
RESD(-1)^2	-0.013266	0.000340	-39.06671	0.0000
F-statistic 267.1(0.000)				

Source: Authors' estimation results with Eviews 10.

In Table 8, the ARCH results for Nigeria show Cambodia's one-period lag of *NEXC* has the coefficient of 0.995107, which is significantly high with a z-statistic of 121.8662. This suggests that the current value of real effective exchange rate is almost entirely determined by its value in the previous period, highlighting strong continuity and minimal variation from historical levels in the short term. The intercept (C) in the mean equation is 0.095810, which is not statistically significant (p=0.5311). This suggests that there are no significant mean changes in the real effective exchange rate independent of its past values, reinforcing the idea that the exchange rate's movements are predominantly influenced by its own inertia. In the variance equation, the constant term (C) is 0.215612, demonstrating a significant level of baseline volatility (z-statistic of 42.92938). This indicates a relatively high inherent volatility in the exchange rate, potentially reflecting the economic fluctuations, policy changes, or market uncertainties prevalent within Cambodia.

Table 9. ARCH results for Myanmar

Mean Equation				
Variable	Coefficient	Std. Error	z-Statistic	Prob.
C	0.433092	0.072750	5.953181	0.0000
NEXC(-1)	0.965299	0.005570	173.3036	0.0000
Variance Equation				
C	0.066289	0.003058	21.67919	0.0000
RESD(-1)^2	2.600196	0.909884	2.857724	0.0043
F=25461.3(0.000)				

Source: Authors' estimation results with Eviews 10.

Table 9 shows nominal exchange rate coefficient in Myanmar is 0.965299. The coefficient is highly significant with a z-statistic of 173.3036, showing that the nominal exchange rate is strongly influenced by its previous value. This coefficient nearly reaching one suggests that the exchange rate in Myanmar demonstrates substantial persistence, meaning that current nominal exchange rate values are almost a direct reflection of the previous period's values. This level of persistence can indicate stability in the currency but might also reflect a rigidity that could impede rapid adjustment to new economic conditions. The variance equation reveals the model's approach to handling volatility. The constant term (C) is 0.066289, with a very high significance level (z-statistic of 21.67919), indicating a relatively moderate baseline volatility in the exchange rate. This finding suggests that while there are fluctuations, they are not excessively volatile under normal conditions, which is favorable for economic planning and foreign trade negotiations.

Table 10. ARCH results for Malaysia

Mean Equation				
Variable	Coefficient	Std. Error	z-Statistic	Prob.
C	0.065025	0.025872	2.513322	0.0120
NEXC(-1)	0.986980	0.004727	208.7983	0.0000
Variance Equation				
C	0.010352	0.000206	50.13430	0.0000
RESD(-1)^2	-0.010740	0.000218	-49.29567	0.0000
F-statistic=1468.92(0.000)				

Source: Authors' estimation results with Eviews 10.

From the ARCH in Table 10 above, the coefficient for one-period lag of real effective exchange rate is remarkably high at 0.986980, with an exceptionally significant z-statistic of 208.7983. This result indicates a very strong persistence in the exchange rate, suggesting that the current *nexc* is almost entirely predictable by its immediate past value. Such a high level of persistence reflects a stable exchange rate environment, where changes from one period to the next are minimal and largely anticipated. The constant term (C) in the mean equation is 0.065025, which is statistically significant (p=0.0120). This signifies that there is a small but consistent adjustment to the real effective exchange rate independent of its previous value, possibly reflecting systematic influences such as policy adjustments or long-term economic trends. In the variance equation, the constant term (C) is 0.010352, indicating a baseline level of volatility that is significant and consistent (z-statistic of 50.13430). This suggests that the underlying volatility of Malaysia's nominal exchange rate is moderate but persistent, providing a foundational level of exchange rate fluctuation that might be attributed to regular market dynamics or external economic influences.

Table 11. GARCH results for Indonesia

Mean Equation				
Variable	Coefficient	Std. Error	z-Statistic	Prob.
C	0.745908	0.567637	1.314058	0.1888
NEXC(-1)	0.982474	0.012734	77.15495	0.0000
Variance Equation				
C	1.661573	0.066105	25.13538	0.0000
RESD(-1)^2	-0.027529	0.002027	-13.57907	0.0000
GARCH(-1)	0.564221	0.003347	168.5564	0.0000

Source: Authors' estimation results with Eviews 10.

The results of the GARCH in Table 11 shows that Indonesia nominal exchange rate with a coefficient of 0.982474, which is highly significant (z-statistic of 77.15495) indicates a very strong autoregressive characteristic, where the current real effective exchange rate is almost completely determined by its value in the previous period. This high degree of persistence suggests that the nominal exchange rate in Indonesia changes gradually over time, providing a predictable pattern based on historical values. The variance equation in the GARCH model is designed to capture the volatility of real effective exchange rate, including terms for both the RESID(-1)^2 and the lagged conditional variance (GARCH(-1)). The coefficient for RESID(-1)^2 is -0.027529, significant with a negative sign (z-statistic of -13.57907), indicating a mean-reversion of volatility. This implies that higher volatility in one period tends to be followed by reduced volatility, aligning with typical financial time series behavior where high-volatility events often stabilize over time. The GARCH(-1) coefficient is 0.564221, significantly positive (z-statistic of 168.5564), highlighting the persistence of volatility. This means that if the exchange rate was volatile in the past, it is likely to remain volatile, pointing to a sustained impact of past volatility on current volatility levels.

Table 12. GARCH results for Singapore

Mean Equation				
Variable	Coefficient	Std. Error	z-Statistic	Prob.
C	0.355531	0.176263	2.017048	0.0437
NEXC(-1)	0.970111	0.012316	78.76919	0.0000
Variance Equation				
C	0.086710	0.037361	2.320868	0.0203
RESD(-1)^2	-0.015018	0.001411	-10.64321	0.0000
GARCH(-1)	0.864199	0.063414	13.62782	0.0000

Source: Authors' estimation results with Eviews 10.

In Table 12, the ARCH and the GARCH coefficients for Singapore nominal exchange rate shows the significant volatility its economy has experienced due to various economic sanctions and commodity price fluctuations. The one-period lagged value of exchange rate is 0.970111, which is highly significant (z-statistic of 78.76919). This suggests that the exchange rate from one period strongly influences the rate in the subsequent period, indicating a gradual adjustment to new information and a tendency for the exchange rate to follow a smooth path over time. In the variance equation, the baseline volatility of the real effective exchange rate is captured by the constant term (C=0.086710), which is significant (p=0.0203). This suggests a foundational level of volatility inherent in the Singapore exchange rate market is influenced by economic policy uncertainty. The GARCH(-1) coefficient of 0.864199 is significantly positive, indicating that past volatility has a strong predictive power on future volatility. This high persistence in volatility suggests that shocks to the exchange rate have long-lasting effects, which could exacerbate the impact of political events on the market stability.

Table 13. GARCH results for Philippines

Mean Equation				
Variable	Coefficient	Std. Error	z-Statistic	Prob.
C	0.173271	0.192182	0.901599	0.3673
NEXC(-1)	0.979277	0.017725	55.24747	0.0000
Variance Equation				
C	0.037690	0.043979	0.857004	0.3914
RESD(-1)^2	-0.006025	0.001299	-4.639478	0.0000
GARCH(-1)	0.561294	0.512449	1.095317	0.2734

Source: Authors' estimation results with Eviews 10.

The GARCH results in Table 13 suggest that the current nominal exchange rate is heavily influenced by its immediate past value, reflecting a stable and predictable behaviour in the short term. This level of persistence is typical for economies with stable macroeconomic policies where the exchange rate adjusts gradually to changes. The intercept (C) of 0.173271, although not statistically significant (p=0.3673), indicates a minor constant impact on the real effective exchange rate that is not explained by its historical performance. The GARCH(-1) coefficient of 0.561294, although not significant (p=0.2734), suggests some degree of volatility persistence. This indicates that while past volatility influences current volatility, the effect is not as strong as it might be in more volatile or unstable economies. For India, the GARCH model's shows how past exchange rate levels and their volatility affect current and future rates.

The results of the GARCH result in Table 14 show that nominal exchange rate shows a high coefficient of 0.976900 with a significant z-statistic of 45.54392. This indicates that the nominal exchange rate is heavily influenced by its previous values, showcasing strong persistence. Such a high level of autoregression suggests that changes in the nominal exchange rate are gradual and predictable over short intervals, typical for an economy where exchange rates are managed within a policy framework designed to maintain stability. The intercept (C) of 0.120945, though not significant (p=0.3652), indicates a minimal baseline impact on real effective exchange rate, which is not explained merely by its lagged values. This denotes underlying macroeconomic trends or policy shifts that exert a constant but subtle influence on the real effective exchange rate. The GARCH(-1) coefficient is 0.565968, which is not significant (p=0.1795). This suggests that past volatility has a moderate influence on future volatility, this relationship is not as strong as might be seen in more freely floating currencies.

Table 14. GARCH results for Thailand

Mean Equation				
Variable	Coefficient	Std. Error	z-Statistic	Prob.
C	0.120945	0.133560	0.905547	0.3652
NEXC(-1)	0.976900	0.021450	45.54392	0.0000
Variance Equation				
C	0.016618	0.016093	1.032599	0.3018
RESD(-1)^2	-0.008602	0.000392	-21.95920	0.0000
GARCH(-1)	0.565968	0.421639	1.342305	0.1795

Source: Authors' estimation results with Eviews 10.

Table 15. GARCH results for Vietnam

Mean Equation				
Variable	Coefficient	Std. Error	z-Statistic	Prob.
C	0.355531	0.176263	2.017048	0.0437
NEXC(-1)	0.970111	0.012316	78.76919	0.0000
Variance Equation				
C	0.086710	0.037361	2.320868	0.0203
RESD(-1)^2	-0.015018	0.001411	-10.64321	0.0000
GARCH(-1)	0.864199	0.063414	13.62782	0.0000

Source: Authors' estimation results with Eviews 10.

The GARCH model results of Table 15 above for Vietnam's *NEXC* lag one period shows a coefficient of 0.970111 and a z-statistic of 78.76919, indicating extremely high statistical significance. This strong autoregressive component suggests that the *NEXC* is highly persistent, with past values being a strong predictor of future rates. Such behavior indicates a stable but slowly adjusting exchange rate environment, where changes are gradual and not abrupt, likely reflecting Vietnam's monetary policy aimed at stabilizing the exchange rate to avoid economic shocks. The variance equation of the GARCH model shows the constant term of 0.086710, significant at the 0.0203 level, indicates a foundational level of volatility inherent to the real effective exchange rate. This level of baseline volatility is influenced by external economic pressures, internal economic policy changes, or market perceptions affecting the South African economy. The GARCH(-1) term at 0.864199, significantly high (z-statistic of 13.62782), shows that past volatility has a strong and persistent influence on current volatility, indicating that volatility shocks tend to have long-lasting effects.

Table 16. GARCH results for Cambodia

Mean Equation				
Variable	Coefficient	Std. Error	z-Statistic	Prob.
C	0.070835	0.157886	0.448649	0.6537
NEXC(-1)	0.996306	0.008428	118.2096	0.0000
Variance Equation				
C	0.074293	0.013774	5.393590	0.0000
RESD(-1)^2	-0.013525	0.002767	-4.888788	0.0000
GARCH(-1)	0.663125	0.062997	10.52629	0.0000

Source: Authors' estimation results with Eviews 10.

Table 16 reported GARCH model analysis of Nigeria's nominal exchange rate. The significant autoregressive component in the mean equation, with nominal exchange rate showing a coefficient of 0.996306 and an exceptionally high z-statistic of 118.2096, demonstrates extreme persistence. This implies that the *NEXC* is almost perfectly predicted by its value in the preceding period, suggesting that the exchange rate evolves in a highly predictable manner with little deviation from its historical path. This can be indicative of a tightly managed exchange rate system where policy interventions ensure stability and reduce unpredictability in the forex market. The GARCH(-1) term at 0.663125 (significant with a z-statistic of 10.52629) indicates that previous periods' volatility has a substantial carryover effect into current volatility. This points to a scenario where shocks to the exchange rate can have prolonged impacts, influencing future volatility levels significantly.

Table 17. GARCH results for Myanmar

Mean Equation				
Variable	Coefficient	Std. Error	z-Statistic	Prob.
C	0.193976	0.162216	1.195786	0.2318
NEXC(-1)	0.988742	0.009222	107.2133	0.0000
Variance Equation				
C	0.049900	0.041271	1.209101	0.2266
RESD(-1)^2	-0.012306	0.003290	-3.740106	0.0002
GARCH(-1)	0.800592	0.168963	4.738260	0.0000

Source: Authors' estimation results with Eviews 10.

Table 17 reported GARCH results and shows that the coefficient of nominal exchange rate lagged one-period is 0.988742. It is highly significant (z-statistic of 107.2133). This indicates that the *NEXC* is predominantly influenced by its value in the previous period, signifying strong continuity and predictability in exchange rate movements. Such high persistence often characterizes exchange rate systems where policy interventions aim to maintain stability or where economic conditions do not fluctuate dramatically in the short term. The GARCH(-1) term at 0.800592, with a significant z-statistic of 4.738260, reflects high volatility persistence. This finding suggests that once the nominal exchange rate exhibits volatility; such fluctuations are likely to continue into the future, emphasizing the impact of past volatility on current and future volatility levels.

The results of the GARCH in Table 18 for Malaysia show that the coefficient of one-period lagged nominal exchange rate stands at 0.986980 with a strikingly high z-statistic of 136.7875, indicating an extremely high level of persistence. This suggests that the current nominal exchange rate is almost entirely dependent on its immediate past value, reflecting a stable and predictable exchange rate behavior over the short term. This characteristic is often indicative of an economy with effective regulatory oversight where exchange rates are managed to ensure gradual adjustments rather than abrupt swings, providing stability which is crucial for international trade and investment planning. The GARCH(-1) coefficient of 0.566803, significant at the 0.0000 level, illustrates substantial volatility persistence. This indicates that volatility in one period tends to influence volatility in subsequent periods, pointing to the impact of economic news and shocks having a prolonged effect on exchange rate stability.

Table 18. GARCH results for Malaysia

Mean Equation				
Variable	Coefficient	Std. Error	z-Statistic	Prob.
C	0.064700	0.039519	1.637168	0.1016
NEXC(-1)	0.986980	0.007215	136.7875	0.0000
Variance Equation				
C	0.006818	0.001503	4.535359	0.0000
RESD(-1)^2	-0.016333	0.003863	-4.227969	0.0000
GARCH(-1)	0.566803	0.094776	5.980442	0.0000

Source: Authors' estimation results with Eviews 10.

Table 19. FIGARCH-DCC results for Indonesia

Mean Equation				
Variable	Coefficient	Std. Error	z-Statistic	Prob.
C	0.541758	0.006526	83.02131	0.0000
NEXC (-1)	0.988435	3.42E-05	28875.60	0.0000
Variance Equation				
C	0.912624	0.072890	12.52051	0.0000
ARCH (-1)^2	-0.150979	0.066743	-2.262107	0.0237
GARCH	0.466768	0.045760	10.20043	0.0000
d-coefficient	0.984710	0.001349	729.9555	0.0000

Source: Authors' estimation results with Eviews 10.

The FIGARCH analysis in Table 19 for Indonesia shows that this model is particularly suitable for capturing long-memory features in the volatility, indicating that shocks to the exchange rate can have persistent effects over time. The coefficient of 0.541758, significantly different from zero (p-value=0.0000), suggests a robust long-term average level around which REER oscillates. This magnitude shows a stable central tendency in nominal exchange rate fluctuations over the long run. A coefficient of 0.988435 with an extremely low standard error and zero p-value implies a high level of persistence in the nominal exchange rate. This suggests that the exchange rate tends to follow a path-dependent process where past values are a strong predictor of current values, reflecting a slow adjustment to new equilibriums. In the variance equation interpretation, the coefficient of constant is 0.912624 signifies the baseline level of volatility when other effects are absent. The significance (p=0.0000) indicates that this base level of volatility is reliably different from zero, emphasizing inherent volatility in the exchange rate. The ARCH coefficient of -1.499766 suggests a negative impact of past squared residuals on current volatility, the GARCH term (long memory) showed the value of 0.466768 confirms the persistence of volatility over time, with significant long-term memory (p=0.0000). This implies that once the exchange rate becomes volatile, this condition is likely to persist, influencing future volatility levels significantly. These results show the complex and persistent nature of volatility in Indonesia's exchange rate. They suggest that any shocks to the exchange rate, whether positive or negative, have long-lasting effects on future volatility levels.

Table 20. FIGARCH-DCC results for Singapore

Mean Equation				
Variable	Coefficient	Std. Error	z-Statistic	Prob.
C	0.346101	0.002014	171.8881	0.0000
NEXC(-1)	0.978616	5.63E-05	17396.15	0.0000
	Variance Equation			
C	0.022324	0.008251	2.705445	0.0068
ARCH(-1)^2	-0.187161	0.007678	-24.37756	0.0000
GARCH	0.697970	0.001073	650.7496	0.0000
d-coefficient	0.563809	0.032126	17.549928	0.0000

Source: Authors' estimation results with Eviews 10.

The FIGARCH results for Singapore in Table 20 shows that the highly significant constant term in the REER suggests a stable long-term average level for the exchange rate in the absence of new information affecting the market. The coefficient of the lagged nominal exchange rate is close to 1, indicating that past values of the *nexc* are a strong predictor of its current level. This persistence characteristic suggests that the *nexc* is highly influenced by its historical values. The constant in variance equation captures the baseline level of variance (volatility) when other terms are zero. A positive and significant indicates that there is a base level of volatility in the real effective exchange rate that is not explained by the model's other terms. ARCH Term showed negative coefficient for the ARCH term, which measures the impact of the squared residuals, is significant and suggests a negative relationship between past shocks and current volatility. This could imply a mean-reverting behavior in volatility, which might indicate that high volatility is likely to be followed by reduced volatility and vice versa. The highly significant and positive coefficient of the GARCH term confirms that volatility shocks have a long-lasting effect on future volatility. This long memory component implies that large changes in volatility can influence the nominal exchange rate volatility for a prolonged period. The presence of a long memory in volatility suggests that the Singapore foreign exchange market experiences persistent effects from shocks, which might be related to economic sanctions, oil price fluctuations, or other structural changes in the economy.

Table 21. FIGARCH-DCC results for Philippines

Mean Equation				
Variable	Coefficient	Std. Error	z-Statistic	Prob.
C	0.139271	3.36E-06	41432.90	0.0000
NEXC(-1)	0.983293	1.30E-07	7546266.	0.0000
	Variance Equation			
C	-1.235696	0.127089	-9.723098	0.0000
ARCH(-1)^2	-0.472782	0.055149	-8.572766	0.0000
GARCH	0.558923	0.040396	13.836150	0.0000
d-coefficient	0.678496	0.00894	75.894407	0.0000

Source: Authors' estimation results with Eviews 10.

The FIGARCH model results of Table 21 for Philippine's nominal exchange rate shows that for real effective exchange rate, signaling a high degree of persistence in the exchange rate movements. This suggests that the nominal exchange rate tends to follow a path dependent on its historical values, which is a common characteristic of financial time series data. Moving to the variance equation, The ARCH term, representing the short-term impact of past squared residuals on current volatility, is negative, which is contrary to usual findings where past volatility spikes increase current volatility. This negative value might reflect a market anomaly or an unusual market reaction to volatility where an increase in past volatility leads to a decrease in current volatility, possibly due to interventions or regulatory actions that aim to stabilize the market. The GARCH term is positive and less than one, suggesting that while shocks to volatility are persistent, they do not have a permanent effect. This term indicates the long-term persistence of volatility, showing that volatility shocks decay over time, but at a rate that is slower than in a standard GARCH model, which is indicative of the long memory feature of the FIGARCH model. The model suggests that while the nominal exchange rate is stable in the short term, there is a significant reaction to shocks, and these shocks have long-term effects in Philippines.

Table 22. FIGARCH-DCC results for Thailand

Mean Equation				
Variable	Coefficient	Std. Error	z-Statistic	Prob.
C	0.127196	2.3E-105	5.6E+103	0.0000
NEXC(-1)	0.971031	1.7E-104	5.9E+103	0.0000
	Variance Equation			
C	-2.149647	0.030939	-69.48056	0.0000
ARCH(-1)^2	3.114303	0.047661	65.34264	0.0000
GARCH	0.805219	0.001091	737.8327	0.0000
d-coefficient	0.95863	0.000067	14307.910	0.0000

Source: Authors' estimation results with Eviews 10.

The results of Table 22 show a complex dynamic structure of volatility related to past information and shocks. The coefficient C is significant given the reported z-statistic and probability, suggests a strong persistent component in the volatility, meaning that shocks to the exchange rate have long-lasting effects. The high z-statistic point out that the coefficients are significantly different from zero, providing strong evidence of the predictive power of the model. The value of the real exchange rate has a coefficient very close to one, which signifies that the exchange rate follows a near-random walk behavior, with past values being highly predictive of current values. The ARCH term, is positive and significantly different from zero, suggesting that past squared shocks have a direct and proportionate impact on current volatility, indicating a typical ARCH effect. GARCH term, is positive and significantly different from zero, pointing to the persistence of volatility shocks over time. This long memory component suggests that volatility can be forecasted using historical volatility data due to its persistent nature. For Thailand, this model's results imply that the exchange rate is influenced by its own past movements and that shocks to the exchange rate persist over time, with negative shocks potentially having a more significant impact on future volatility.

Table 23. FIGARCH-DCC results for Vietnam

Mean Equation				
Variable	Coefficient	Std. Error	z-Statistic	Prob.
C	0.346101	0.002014	171.8881	0.0000
NEXC(-1)	0.978616	5.63E-05	17396.15	0.0000
	Variance Equation			
C	0.022324	0.008251	2.705445	0.0068
ARCH (-1)^2	0.187161	0.007678	24.37756	0.0000
GARCH(-1)	0.697970	0.001073	650.7496	0.0000
d-coefficient	-0.771274	0.014928	-51.66757	0.0000

Source: Authors' estimation results with Eviews 10.

In Table 23, the constant coefficient represents the long-term average of the log-volatility process, which is statistically significant with a z-statistic of 171.8881 and a p-value of 0.0000, indicating a highly reliable estimator of the average volatility of real effective exchange rate over the sample period. The coefficient of RER(-1) near one with a very small standard error suggests that past values of real exchange rate heavily influence its current level, indicating a high level of persistence of exchange rate in Vietnam. The ARCH reflects the immediate impact of market shocks on volatility, which is negative and highly significant, suggesting a pronounced reaction of volatility to recent market changes. This point towards a heavy-tailed distribution of returns where large swings in real effective exchange rate are more common than a normal distribution would predict. The GARCH long memory component, is significantly different from zero, suggesting that shocks to the exchange rate volatility have a persistent effect that decays very slowly over time. However, for Vietnam, this mean that the exchange rate is influenced by its past values and reacts significantly to new information (represented by the ARCH term), particularly to negative market events. The presence of a long memory component suggests that these effects are not short-lived but continue to influence volatility over time. The FIGARCH results of Table 23 also allows for asymmetry in the volatility response to positive and negative shocks.

Table 24. FIGARCH-DCC results for Cambodia

Mean Equation				
Variable	Coefficient	Std. Error	z-Statistic	Prob.
C	0.000610	1.48E-05	41.15408	0.0000
NEXC(-1)	0.999964	7.83E-05	12774.77	0.0000
	Variance Equation			
C	-3.127534	2.099758	-1.489474	0.1364
ARCH(-1)^2	0.006364	0.005159	1.233598	0.2174
GARCH	0.611321	0.267444	2.285788	0.0223
d-coefficient	0.542930	0.001925	282.04155	0.0000

Source: Authors' estimation results with Eviews 10.

In Table 24, the constant (C) has a near-zero coefficient (0.000610) with a highly significant z-statistic. This indicates statistically significant intercept in the conditional mean equation for *NEXC*. This suggests that nearly all the predictability in nominal exchange rate can be accounted for by its own past values. The coefficient very close to one (0.999964) with a highly significant z-statistic shows a strong persistence of nominal exchange rate, indicating that the past value is a nearly perfect predictor of its current value, a characteristic of a random walk. Coefficient of ARCH is positive but not statistically significant (p=0.1319), suggesting that recent shocks have a minimal and not statistically significant impact on current volatility. The GARCH indicates the long-term persistence of volatility (GARCH effect), with a coefficient that is statistically significant (p=0.0223), indicating that volatility shocks have a persistent impact over time. The analysis suggests that the nominal exchange rate is characterized by a high level of persistence and potential heavy tails in its distribution, but recent shocks and their asymmetric impact are not playing a significant role in forecasting current volatility.

Table 25. FIGARCH-DCC results for Myanmar

Mean Equation				
Variable	Coefficient	Std. Error	z-Statistic	Prob.
C	0.136302	0.002126	64.12275	0.0000
NEXC(-1)	0.993281	3.02E-05	32924.95	0.0000
	Variance Equation			
C	-0.187750	0.006239	-30.09357	0.0000
ARCH (-1)^2	-0.321892	0.041477	-7.760684	0.0000
GARCH (1)	0.579298	0.012131	47.75296	0.0000
d-coefficient	0.808820	0.002469	327.59011	0.0000

Source: Authors' estimation results with Eviews 10

In Table 25 above, the coefficient C, representing the intercept in the variance equation, is highly significant with a near-zero probability (p-value), indicating a strong baseline level of volatility in the exchange rate movements. The lagged variable of nominal exchange rate RER(-1) with a very high z-statistic and a p-value of zero confirms the past nominal exchange rate values are extremely predictive of current *NEXC* values, signifying a strong autoregressive character in the exchange rate series. The coefficients ARCH term and its residual represent the short-run components of volatility, the ARCH term, and the measure of asymmetry in the impact of residuals (leverage effect), respectively. The negative sign of ARCH, along with its significant z-statistic, suggests a less than proportional reaction of volatility to past squared residuals, The coefficient of GARCH term and its high z-statistic and zero probability reveal that past volatility is highly predictive of future volatility, signifying a long memory characteristic of the volatility process. This could mean that volatility shocks to the nominal exchange rate are not only impactful in the short run but also persist over a longer period, influencing future volatility levels. Thus, in Myanmar, an emerging economy with active engagement in bitcoin and other cryptocurrencies, this volatility dynamic could be crucial. A higher persistence in volatility indicates that external shocks, possibly including those related to bitcoin demand fluctuations, have long-lasting effects on the stability of the real effective exchange rate. Such findings reflects the impact decisions related to risk management, investment, and economic policy, as sustained volatility in exchange rates may affect international trade, investment flows, and economic stability.

The FIGARCH for Malaysia are reported in Table 26. The FIGARCH model results for Malaysia's nominal exchange rate signify a complex dynamic of volatility in exchange rate movements influenced by Bitcoin demand and other economic factors. The constant being negative and significant indicates that the baseline volatility is subject to mean reversion. The ARCH coefficient is also negative and significant; suggesting that past volatility has a strong shock impact on current volatility. This can be indicative of a market that reacts strongly to past volatility shocks, where the effects of large changes tend to be followed by further large changes, which may remain over long periods. This trait is particularly important for financial risk management as it points to a potentially higher risk environment for investors and policymakers. The GARCH coefficient being positive and significant suggests that volatility shocks are persistent over time, which aligns with the concept of long memory in volatility. This indicates that the effects of shocks to volatility do not decay quickly and that past periods of instability can influence future volatility over a long horizon. The FIGARCH model indicates a complex volatility structure in real effective exchange rate, likely driven by both external economic forces and internal policy measures.

Table 26. FIGARCH-DCC results for Malaysia

Mean Equation				
Variable	Coefficient	Std. Error	z-Statistic	Prob.
C	0.004035	4.41E-05	91.56762	0.0000
NEXC(-1)	0.999141	1.23E-05	81386.46	0.0000
Variance Equation				
C	-3.216785	0.157172	-20.46668	0.0000
ARCH(-1)^2	-0.406821	0.014469	-28.11642	0.0000
GARCH (-1)	0.250863	0.036499	6.873086	0.0000
d-coefficient	0.452890	0.001167	388.08054	0.0000

Source: Authors' estimation results with Eviews 10.

Discussion

When comparing our research findings with those from other studies, it is important to note that while Yuliadi et al. (2024) confirm low volatility in currency exchange rates due to Singapore’s managed floating exchange rate system, our research findings which show a significant persistence of shocks in the exchange rate of the Singaporean currency do not entirely agree with these findings. Our research findings indeed corroborated those of Yuliadi et al. (2024), who indicated that the government should keep an eye out for economic measures that would lessen the impact of ASEAN countries’ exchange rate volatility.

In addition, our results are consistent with those of Rossanto et al. (2023), who found that currency exchange rate volatility in nations that adhered to the free float regime persisted even ten years after the financial crisis; Jonathan (2022) obtained evidence of an asymmetric response to exchange rate fluctuations that corroborates ours; Ain (2022) produced results from the wavelet power spectrum analysis that are consistent with our results for ASEAN nations. According to Ain’s (2022) research findings; four ASEAN countries namely, Singapore, Thailand, Malaysia, and Indonesia were found to have highly volatile currency rates. Thailand has little short-term volatility and no increased long-term volatility, in contrast to the Philippines’ slight volatility. Our study’s outcome suggests that the real exchange rate in Philippines exhibits time-varying volatility that can be partially captured by its own past values. This research outcome is consistent with those of Abdul, Hsia Hua Sheng, and Natalia Diniz-Maganini (2021). These authors demonstrated that exchange rate volatility is asymmetric and time-varying by using empirical research based on the EGARCH (1, 1) specification fitted on monthly Asian currencies. The findings indicate that three currencies have indications of asymmetry in their conditional variance prior to the Asian crisis. With the exception of one, all of them displayed a noticeable increase in volatility and asymmetry effect.

The findings from our study also corroborate those reported by Goda & Priewe (2020). In fact, the authors evaluated the primary causes of cyclical REER behavior in emerging market economies (EME). The research employed a sample size of fifteen emerging market economies from 1996 to 2016, covering several significant economic shocks such as the Asian crises of 1997–1998; the Russian crisis of 1998;

and the Turkish crisis of 2001. The 15 nations are divided into two categories by the study: industrial emerging market economies and commodities developing market economies. The dynamic panel fixed effects model was the econometric method employed to determine the factors influencing REER. Based on the findings, developing market economies that rely on commodities typically have higher REER volatility.

5. Conclusion

The study evaluated the behavior of the nominal exchange rate in ASEAN countries, namely, Indonesia, Singapore, the Philippines, Thailand, Vietnam, Cambodia, Myanmar, and Malaysia. The current research aims to analyse the behavior of exchange rates in ASEAN countries using quantile regression sensitivity analysis. ARCH, GARCH, and FIGARCH modeling are also performed for robustness checks. This study made use of daily exchange rate data covering the period from January 1, 1990, to December 30, 2023, inclusive of weekends and holidays. The data were handled in two eras (before and after the financial crisis). The results are discussed based on each country. The models are well explained, and the results are presented in an organized manner. The GARCH and FIGARCH-DCC modeling techniques were executed to further gauge the impulsive performance of exchange rates of the aforementioned countries. Our research findings uphold asymmetry and persistence in the behavior of exchange rates of ASEAN currencies. In terms of comparative discussions, our research findings considerably align with the findings of other researchers, namely Yuliadi et al. (2024), Rossanto et al. (2023), Jonathan (2022), Ain (2022), and Natalia et al. (2021).

In Indonesia, the FIGARCH results show high persistence in volatility, as indicated by a significant GARCH term, suggesting that shocks to the exchange rate have lasting effects. Singapore's exchange rate exhibited significant persistence of shocks, indicating that the effects of shocks on the exchange rate are enduring. Philippines's FIGARCH model illustrates a high level of volatility persistence. For Thailand, the volatility is characterized by long-memory and a considerable volatility shock, which reflects the sensitivity of the exchange rate to market dynamics. In Vietnam, exchange rate volatility does not die off quickly, and this has the tendency to stimulate speculative chances. Cambodia exhibits weighty persistence in exchange rate volatility. Myanmar's FIGARCH analysis shows permanent exchange rate shocks. Lastly, in Malaysia, volatility persistence in the nominal exchange rate is robust. It suffices to advise that all ASEAN governments should enhance regulatory frameworks that monitor and possibly control exchange rate transactions to prevent excessive speculative activities that could destabilize national currencies. Accordingly, we recommend the need for ASEAN countries' central banks to put into effect the preventative measures necessary to preserve economic stability and control the volatility dynamics of the foreign exchange market.

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The impact of China's foreign trade on their Actual-Open Emissions of CO_2 in the years 2000–2020 in the context of EU energy policy

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Aim: This article aims to analyze the impact of China's trade with 78 major trading partners on Actual-Open Emission of CO_2 (EAO) from 2000 to 2020 in light of the European Union's (EU) goal to reduce CO_2 emissions by 20% by 2020 compared to 1990 levels.

Research Methods: The research is based on the Actual-Open Emission of CO_2 model and employs the circular flow model to assess the influence of China's foreign trade on CO_2 emissions during the years 2000–2020.

Findings: The study revealed that China's foreign trade significantly influenced its CO_2 emissions in all years analyzed, with positive contributions to EAO due to a trade surplus (exports exceeding imports). As the world's largest exporter and the second-largest importer, China's trade activity resulted in substantial CO_2 emissions. Four key indicators were identified as influencing the difference between Official-Close Emission of CO_2 (EOC) and EAO: China's GDP, the percentage of exported GDP, the percentage of imported GDP, and EOC levels. These findings highlight the significant role of trade in China's CO_2 emissions, which is critical in the context of EU initiatives like "Fit for 55."

Keywords: International trade, China, CO_2 emissions, EU energy policy.
JEL: F18, Q54, Q56,

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1. Introduction

The second biggest economy in the world in the years 2000–2020 was China (worlddata.info). They were in first place in terms of GDP PPP from the year 2014 (Gentle 2016: 87). From 2000 to 2020, China was among the three countries with the world's most prominent export and importers (unctad.org; wits.worldbank.org). China's trade significantly impacted the natural environment, including CO_2 emissions. The considerations contained in this study result from the energy policy implemented by the EU. One of its basic assumptions in 2007–2020 was the rule – 3 times 20%.

EU policy and its energy policy directly refer to sustainable development (SD). Gro Harlem Brundtland proposes the basic definition of SD presented in the report "Our Common Future." SD is defined here as: "meeting the needs of the present without compromising the ability of future generations to meet their own needs." (Czaja, Becla 2002: 308–309; Górka et al. 1995: 78; Rao 2000: 85; Adamczyk 2001: 28–29).

We live in a global world, and therefore, the activities of particular countries have a direct or indirect impact on others. Nevertheless, it does not mean that all countries function in the same way and follow the same rules. Some countries contribute significantly to reducing global CO_2 emissions, bearing high costs compared to others; nonetheless, it does not bring the intended effects of an absolute reduction of CO_2 emissions. This study confirms the issue, especially regarding China's trade with 78 countries. The struggle of the EU with the issue of CO_2 emissions does not affect China with this problem. This study will show that China had, between 2000 and 2020, an impact on world CO_2 emissions lower than officially shown because they export more CO_2 than imported. By this example, it will also be shown that CO_2 emissions are a global problem, and the struggle of several countries does not change much in the universal aspect. Fundamental questions are as follows.

How significant influence did China trade on CO_2 emission globally? What should the EU do to make its energy policy more efficient? Should not all countries be involved in efforts to reduce CO_2 emissions? Should solutions be undertaken to encourage other countries to take a similar approach to the CO_2

issue? Should the European Union change its approach regarding CO_2 reduction and take international exchange into account?

This new approach should help to answer these questions.

China is one of the world leaders in terms of the value of CO_2 emissions and international exchange. According to the Economic Complexity Index (ECI) (The Observatory), it is the world's second-largest importer and exporter and one of the most complex economies.

The primary purpose of this paper is to show the impact of foreign trade on Actual-Open Emissions of CO_2 (E_{AO}) in China after considering trade with the 78 countries. It is not about the value of Official-Close Emission of CO_2 (E_{OC}) emissions but about its accurate volume in regard to the CO2 transfer both in export and import products. There should also be services that should have been considered in this research.

The survey is based on a circular economic flow model principle that shows money flows through the economy. There are two kinds of this model. Closed – inside the country, and Open – including export and import factors. The same refers to the open and closed economy. The Official-Close Emission of CO_2 is similar to the close circular flow model concept. Is it the right approach to this problem? It seems not to be. We live in a global world where international trade is one of the economy's most important and influential factors. This factor greatly influences CO_2 emission because producing goods and services accompanies CO_2 emission.

2. Methods

From 2007, the EU energy policy was created by rules, mechanisms, and economic and financial instruments (Komunikat UE KOM (2007) 1, Dyrektywa 96/61/WE, Dyrektywa 2001/80/WE, Dyrektywa 2003/87/WE, Dyrektywa 2006/32/WE, Dyrektywa 2009/28/WE, Komunikat UE KOM (2010) 639, Komunikat UE KOM (2008) 781, Komunikat UE KOM (2008) 772, Komunikat UE KOM (2006) 105, Komunikat UE KOM (2008) 13, Komunikat UE KOM (2008) 768). It outlines the basic directions for developing the EU energy sector (Jeżowski 2011). Energy policy came into force in the EU in 2007. The European energy policy aimed to achieve 3 × 20% by 2020. It involves the reduction of CO_2 emissions by 20% in 1990, increasing participation of renewable energy sources in the energy mix to 20%, and improving energy use efficiency by 20% compared to 1990.

It should also be emphasized that the indicated aims are interconnected. The last two goals significantly influence the reduction of CO_2 emissions, which in turn impacts the changes in other objectives of the EU energy policy.

The Actual-Open Emission of CO_2 was determined as the CO_2 emissions of a particular country. It is diminished by emissions in exported goods and services of the country and magnified by emissions imported in products and services from the importing country. It means that the emissions balance of CO_2 should decrease Actual-Open Emissions of CO_2. The following formulas present a method used to calculate Actual-Open Emissions of CO_2 for the China:

$$S_B = \left(\frac{I_m}{GDP}\right)\% \times E_{OCC} - \left(\frac{E_x}{GDP}\right)\% \times E_{OC} \quad (1)$$

$$\boldsymbol{E_{AO} = E_{OC} + S_B} \quad (2)$$

S_B – The balance of CO_2 emissions in the selected country;
E_{OC} – The Official-Closed Emissions of CO_2 in China;
E_{OCC} – Official-Closed Emission of CO_2 in selected country;
Ex – Value of the China export to selected country;
Im –Value of the China import from a particular country;
GDP – The gross domestic product of a selected country;
(Im/GDP)% – part of the GDP of a specific country from with the China imported;
(Ex/GDP)% – part of China's GDP which was exported to a particular country;
(Im/GDP)%*E_{OCC} – Quantity of imported CO_2 in goods from a specific country to China;
(Ex/GDP)%*E_{OC} – Quantity of exported CO_2 from China to the particular state in exported goods;
E_{AO} – Actual-Open Emissions of CO_2 in China.

To show the E_{AO} in a specific country, we need the data of all China trade partners. In this survey, we have 78 main trade partners of China divided into five continents: Africa, Asia, Europe, North America, Oceania, and South America (table 1). Africa is represented by four countries, Asia by 29 countries, and Europe by 33 countries. Four countries represent North America, Oceania 2, and South America by seven countries. These countries and China are among the largest CO_2 emitters in the world.

They were responsible for 96–97% of the world's CO_2 emissions from 2000 to 2020. Due to the multitude of data, the research results in this study will be limited to individual continents, with an indication of the countries that had the most significant impact on CO_2 exchange.

Table 1. Countries participating in the study are divided into individual continents

Continent	Country	Continent	Country	Continent	Country
Africa	Algeria	Asia	Singapore	Europe	Lithuania
Africa	Egypt	Asia	Thailand	Europe	Luxembourg
Africa	Morocco	Asia	Turkmenistan	Europe	Latvia
Africa	South Africa	Asia	Turkey	Europe	North Macedonia
Asia	United Arab Emirates	Asia	Uzbekistan	Europe	Netherlands
Asia	Azerbaijan	Asia	Vietnam	Europe	Norway
Asia	Bangladesh	Asia	Chinese Taipei	Europe	Poland
Asia	China	Europe	Austria	Europe	Portugal
Asia	Cyprus	Europe	Belgium	Europe	Romania
Asia	Hong Kong	Europe	Bulgaria	Europe	Russia
Asia	Indonesia	Europe	Belarus	Europe	Slovakia
Asia	India	Europe	Switzerland	Europe	Slovenia
Asia	Iran	Europe	Czechia	Europe	Sweden
Asia	Iraq	Europe	Germany	Europe	Ukraine
Asia	Israel	Europe	Denmark	North America	Canada
Asia	Japan	Europe	Spain	North America	Mexico
Asia	Kazakhstan	Europe	Estonia	North America	Trinidad and Tobago
Asia	South Korea	Europe	Finland	North America	United States (US)
Asia	Kuwait	Europe	France	Oceania	Australia
Asia	Sri Lanka	Europe	United Kingdom	Oceania	New Zealand
Asia	Malaysia	Europe	Greece	South America	Argentina
Asia	Oman	Europe	Croatia	South America	Brazil
Asia	Pakistan	Europe	Hungary	South America	Chile
Asia	Philippines	Europe	Ireland	South America	Colombia
Asia	Qatar	Europe	Iceland	South America	Ecuador
Asia	Saudi Arabia	Europe	Italy	South America	Peru
				South America	Venezuela

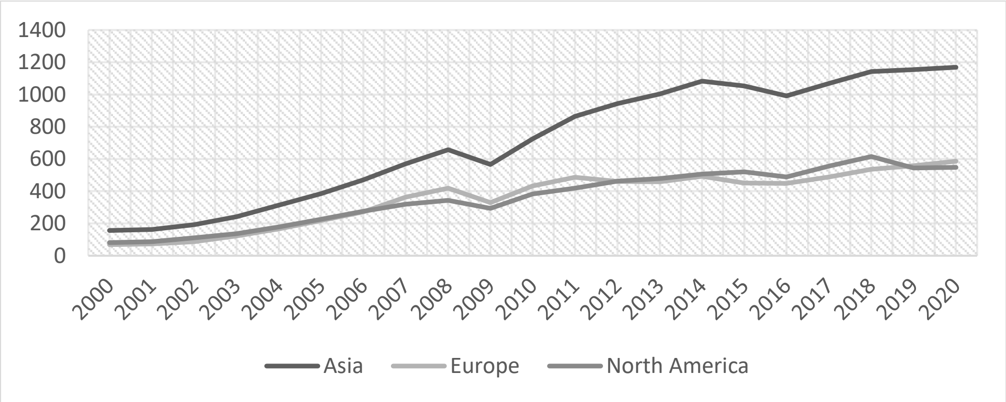
Source: own elaboration.

3. Trade between China and 78 countries from 2000 to 2020

China is the second biggest economy in the world. Total exports of China in years had a grooving trend until 2008, 2009–2014, 2016–2018, and 2019–2020 (figure 1 and figure 2). The declines in total China exports in a survey time were in the years 2009, 2015–2016, and 2019. China's total exports reached 319,71 billion (B) USD in 2000, up to 2491,05 B USD in 2020. It increased almost eight times in the twenty-one years considered in this survey, and by 21 years considered in this survey, China exported a total of 78 countries, 32,35 trillion USD.

Between 2000 and 2020, China's exports to Asia countries achieved value from 156 B of USD in 2000 to 1168 B of USD in 2020. In the case of North American countries, China's exports were between 81 B of USD in 2000 and 615 B of USD in 2018. In survey time, China's exports to Europe reached 69 B USD in 2000 and 586 B in 2020. In 2007–2010 and 2019–2020, China's exports to European countries exceeded North American countries. Between 2000 and 2020, China's exports to South America achieved a value of 4 B USD in 2000 and 93 B USD in 2013. From 2000 to 2020, China's exports to Oceania have yet to reach 65 B USD; they level 43 B USD to Africa.

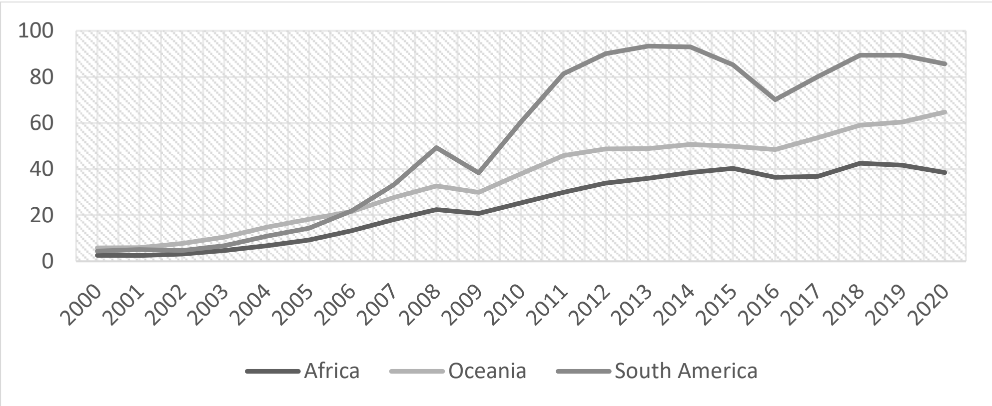
Figure 1. China exports to Asia, Europa and North America in the years 2000–2020 in USD billion



Source: own study based on The Observatory of Economic Complexity.

China's exports to Asia in 2000–2020 constituted between 42.8% (2007) – and 49% (2000) of total China exports, to North America 21.8% (2011) and 27.1% (2002); in the case of Europa, it was 20.5% (2015) and 27.5% (2008). In South America, China exported 1.2% (2002) and 4.4% (2012, 2013) of total their exports. Oceania reached a level of 2.6% and Africa 1.8%. It shows that in international trade, contacts with Asia, North America, and Europe are the most important partners for China.

Figure 2. China exports to Africa, Oceania and South America in years 2000–2020 in USD billion

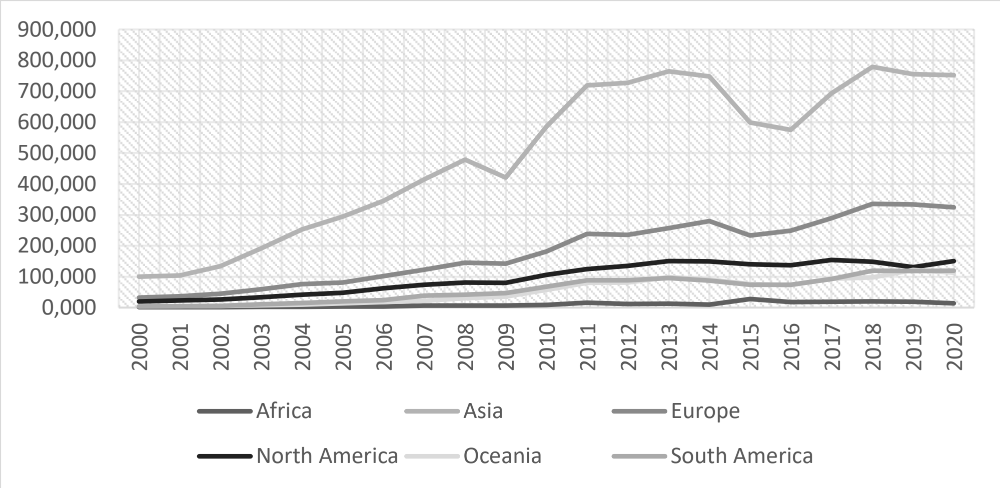


Source: own study based on The Observatory of Economic Complexity.

China's total imports grew in 2000–2008, 2009–2013, and 2016–2018 (figure 3). In the remaining years, the total China imports declined, according to a survey. China's imports reached 160 B USD (2000) to 1499 B USD (2018). It increased almost ten times over the 21 years of the studied period.

Between 2000 and 2020, China's imports from Asia increased from 99.9 B USD in 2000 to 778 B USD in 2018. In the case of European countries, the value of China's imports was between 32 B of USD in 2000 and 335 B of USD in 2018. In survey time, China's imports from North American countries achieved a value of 19 B of USD in 2000 and 1534 B in 2017. Between 2000 and 2020, China's imports from South America reached 3.6 B USD in 2000 and 119 B USD in 2020. China imports from 2000–2020 from Oceania achieved a value of 4.2 B USD in 2000 and 121 B USD in 2029. China's imports from African countries never reached 28 B USD.

Figure 3. China imports from six continents represented by 78 countries in USD billions in 2000–2020



Source: own study based on The Observatory of Economic Complexity.

China imports from the Asian countries in 2000–2020 constituted 63.8% (2005) – 51% (2019 and 2020) of the total China imports. Imports to China from European countries in the survey period were between 17.6% in 2005 and 22.5% in 2019; in the case of North American countries, it was 13.3% (2001) and 8.1% (2019). From South American countries, it was between 2.2% (2000) and 8.1% (2020) of total China imports. China imports from Australia and New Zealand in 2000–2020 constituted 2.4% (2001) – 8.2 (2019) of the total China imports. Africa never reached a level higher than 2.5%. It shows that China's contacts with Asia, North America, and Europe are the most critical partners in imports, similar to China's exports.

From 2000 to 2020, the value of China's imports from Europe fluctuated, but it was generally upward. This time, China had a positive balance in international trade with all European continents except the Island, Germany, and Switzerland. It means that they imported less from them than they exported to them. It means that China is a vital partner for almost all countries because they are massive international suppliers. It shows that in global trade, China was an essential partner for European countries, especially European Union countries. The EU can influence China's CO_2 emission policy to be more restrictive. It can be easier to establish because the trade balance between China and Europe, especially with the EU, was positive. This situation is less challenging to achieve because China has more to lose.

4. Official-Closed Emission of CO_2 of China and the 78 countries

Reducing CO_2 emissions is one of the priorities of the EU energy policy. However, the reduction of CO_2 emissions assumption concerns only the EU, and except for the encouragement, there is no other possibility to convince other countries worldwide to undertake similar actions. Countries participating in this survey were responsible for around 97% of all CO_2 emissions.

Table 2. China Official-Closed Emissions of CO_2 in MT in 1990, 2000–2020

Year	1990	2000	2001	2002	2003	2004	2005	2006
MT of CO2	2323.83	3360.87	3523.08	3843.40	4532.15	5334.89	6098.18	6677.29
China % of World Emission	10.78	14.09	14.57	15.54	17.46	19.54	21.47	22.80
Year		2007	2008	2009	2010	2011	2012	2013
MT of CO2		7239.76	7378.25	7713.90	8145.83	8827.19	9004.24	9247.43
China % of World Emission		23.87	24.11	25.74	26.03	27.44	27.70	27.96
Year		2014	2015	2016	2017	2018	2019	2020
MT of CO2		9293.19	9279.73	9278.98	9466.36	9652.69	9810.46	9899.33
China % of World Emission		28.04	27.95	27.81	28.07	28.10	28.55	30.63

Source: own study based on BP report.

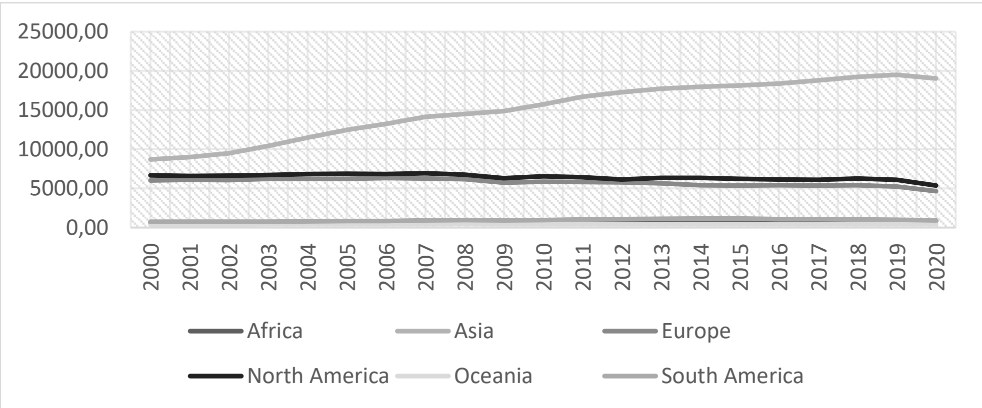
Official-Close Emission of CO_2 is a value of CO_2 emitted by a country's economy. Table 2 and figure 4 presents China, 78 countries, and the whole world's emission of CO_2. From 2004, China was the world's biggest emitter of CO_2, and from 2000–2020, it was responsible for 14.09% (2000) to 30.63% (2020) of the world's emissions. In the survey, China emitted 157.61 BT of CO_2 and recorded 3360.87 – 9899.33 MT of CO_2. In the entire study, China emitted more CO_2 than in 1990.

In the year 1990, the following CO_2 emission values occurred in MT: African countries emitted 510.35, Asian countries – 6175.08, Europa – 7521.44, North

America – 5173.59, Oceania – 300.78 and South America – 535.42. At this time, European countries were the biggest CO_2 emitters in the world. In the years 2000–2020, African countries emitted a total of 16.41 BT of CO_2 and were responsible for around 2.5% of the world's emission of CO_2. In the same period, Asia emitted 316.67 BT of CO_2, equivalent to 36% (2000) and 58.9% (2020) of the world's emission of CO_2. Europa emitted 121.15 BT of CO_2 in survey time and was responsible for 25.2% (2000) – 14.37% of world emission of CO_2. From 2000–2020 North America emitted 134.98 BT of CO_2 and was responsible for 27.93% (2000) – 16.61% (2020) of world emissions. Oceania emitted 8.95 BT, and South America emitted 19.89 BT of CO_2 simultaneously.

The biggest emitter of CO_2 (2000–2020) in North America was the US, and in Asia was China. In the survey, China emitted 157.10 BT of CO_2 and had taken values in the range 9899.33 (2020) – 3360.87 (2000) MT of CO_2. In 2020, China was responsible for 30.6% of the world's CO_2 emissions.

Figure 4. The Official-Closed Emissions of CO_2 on six continents in MT in 2000–2020



Source: own study based on BP report.

If we consider the Official-Closed Emission of CO_2, China did not fulfill in any survey years one of the three main aims of the EU energy policy – reduction of CO_2 emission to the level of 80% emission since the year 1990 (table 3). At this point it is impossible to present if continents fulfill the EU obligation because it must be calculated for each country separately. Moreover, there was a noticeable upward trend in China's Official-Closed Emission of CO_2 throughout the entire period under study. Except for 2015 and 2016, in the whole period under study, the volume of this emission was higher than that recorded in the year 1990.

Table 3. China % of E_{OC} about 1990 in the years 2000–2020

Year	2000	2001	2002	2003	2004	2005	2006
% of E_{OC}	144.63	151.61	165.39	195.03	229.57	262.42	287.34
Year	2007	2008	2009	2010	2011	2012	2013
% of E_{OC}	311.54	317.50	331.95	350.53	379.85	387.47	397.94
Year	2014	2015	2016	2017	2018	2019	2020
% of E_{OC}	399.91	399.33	399.30	407.36	415.38	422.17	425.99

Source: own study based on BP report.

5. Actual-Open Emission of CO_2 – after considering the China trade with the 78 countries

Using the formula (1), the S_B of CO_2 emission was calculated for China and 78 countries from five continents. It shows us the balance of CO_2 emissions. If the value is positive, it means that China's import of CO_2 from a particular country was higher than its export to this country. If the value was negative, then the opposite situation took place – the export of CO_2 was higher than the import. We add the definite number to a particular country's emission amount, and if it is negative, we diminish it. Due to the large number of countries included in the study, individual countries were assigned to appropriate continents. The values given for individual continents are the sum of the CO_2 balance results for the particular countries participating in the study and located on the relevant continent (table 1). Therefore, these data are presented about individual continents. The results of the calculation are presented in the tables 4 and 5.

Table 4. China's balance of CO_2 emissions (S_B) and Actual-Open Emissions of CO_2 (E_{AO}) in 2000–2009 in MT

	2000	2001	2002	2003	2004	2005	2006	2007	2008	2009
Africa	-5.11	-4.54	-5.44	-9.20	-15.31	-20.32	-27.51	-29.78	-27.31	-24.11
Asia	-357.62	-342.52	-398.06	-508.06	-666.85	-819.35	-912.71	-910.95	-774.35	-596.91
Europe	-144.75	-140.82	-182.22	-285.89	-403.75	-536.22	-612.42	-697.46	-626.02	-446.00
North America	-214.61	-214.94	-273.97	-358.49	-467.17	-583.42	-643.76	-622.28	-518.87	-415.34
Oceania	-12.61	-11.44	-15.51	-23.10	-34.62	-41.18	-43.35	-46.22	-40.91	-29.94
South America	-10.26	-10.34	-7.83	-10.95	-21.03	-28.41	-42.85	-52.55	-64.35	-42.33
Total	-744.96	-724.60	-883.03	-1195.69	-1608.74	-2028.90	-2282.61	-2359.24	-2051.81	-1554.64
E_{AO}	2615.91	2798.48	2960.37	3336.46	3726.16	4069.29	4394.68	4880.53	5326.44	6159.26

Source: own study based on BP report, CO_2 EMISSIONS FROM FUEL COMBUSTION Highlights and The Observatory of Economic Complexity.

In the considered period, the total China trade balance of CO_2 emission with the 78 countries taken together was negative, meaning that China imported less CO_2 than exported. China's lowest value in 2007 was -2359.24 MT of CO_2, and the highest was in 2001 -724.60 MT of CO_2. However, if we consider each continent separately, only in the case of African countries, we notice a positive balance of CO_2. It was +1.57 MT of CO_2 (2015). From 2000 to 2020, China had a negative balance with Africa; by those years, it was -318.08 MT. In the case of the rest of the continents, China exported more CO_2 than imported every year considered period. This means that their CO_2 balance was negative throughout the entire research period. Asia was the continent with China having the lowest negative balance of CO_2. Its value ranged from -342.52MT of CO_2 (2001) to -912.71 MT of CO_2 (2006), and by the considered period, it was -11904.86 MT of CO_2. The next continent with which China had a negative balance of CO_2 was North America. The CO_2 balance ranged from -643.76 MT (2006) to -214.61 MT (2000), and in the years 2000–2020, the total value was -8702.02 MT of CO_2.

The next continent with which China had a negative balance of CO_2 was Europa, and the value was between -140.82 MT in 2001 and -697.46 MT in the year 2007. The total balance in the survey time was -7955.21 MT of CO_2. In the case of South America, China had a negative balance of CO_2; its value was between -7.83 MT in 2002 and -71.18 MT in 2011. In all survey years, it was -798.84 MT of CO_2. A similar situation was with Oceania. The CO_2 balance with China was at a value of -11.44 MT in 2001 and -46.22 in 2007. The total value in 2000–2020 was -523.17 MT of CO_2.

Table 5. China's balance of CO_2 emissions (S_B) and Actual-Open Emissions of CO_2 (E_{AO}) in 2010–2020 in MT

	2010	2011	2012	2013	2014	2015	2016	2017	2018	2019	2020
Africa	-25.75	-19.83	-25.02	-22.50	-23.55	1.57	-4.18	-5.75	-8.46	-6.91	-9.07
Asia	-640.22	-629.20	-614.15	-561.80	-558.87	-533.60	-496.84	-445.57	-379.02	-371.54	-386.65
Europe	-519.74	-497.89	-414.36	-372.09	-359.65	-298.11	-284.64	-285.71	-269.37	-280.02	-298.07
North America	-475.05	-447.66	-445.49	-414.88	-403.81	-393.67	-363.59	-385.51	-386.71	-339.12	-333.68
Oceania	-31.36	-31.32	-31.17	-23.30	-21.19	-19.69	-16.45	-14.04	-14.91	-7.57	-13.31
South America	-61.97	-71.18	-70.59	-63.68	-58.95	-47.59	-34.49	-32.35	-23.65	-22.78	-20.71
Total	-1754.09	-1697.08	-1600.78	-1458.23	-1426.02	-1291.08	-1200.20	-1168.92	-1082.12	-1027.94	-1061.49
E_{AO}	6391.74	7130.11	7403.47	7789.19	7867.17	7988.65	8078.78	8297.44	8570.57	8782.51	8837.85

Source: own study based on BP report, CO_2 EMISSIONS FROM FUEL COMBUSTION Highlights and The Observatory of Economic Complexity.

China had the lowest negative S_B of CO_2 emission for 21 years of the survey, with the US (-7367.21 MT), Hong Kong (-5055 MT), Japan (-2938.62 MT), Germany (-1489.23 MT), United Kingdom (-1153.05 MT) and Netherlands (-1016.23 MT). The only countries with which China had a positive value of S_B of CO_2 by the survey years were Uzbekistan (10 MT), Turkmenistan (112 MT), Saudi Arabia (118 MT), Qatar (15 MT), Oman (158 MT), Kuwait (25 MT), Kazakhstan (50 MT), Iraq (45 MT) and Iran (122 MT). In the case of South America: Brazil (-342.35 MT), Oceania: Australia (-433.56 MT) and Africa: Egypt (-140.48 MT).

China had negative S_B of CO_2 emission in all 21 years, as considered in this survey with all EU countries (-7561.69 MT). This means the EU had an E_{AO} higher than the E_{OC} because China exported much more CO_2 to EU countries than it did to China. This situation influenced meeting the requirements of the EU energy policy – 20% less CO_2 emissions compared to 1990. The same is true in the case of E_{OC}.

Table 6. China % of E_{AO} and E_{OC} about 1990 in 2000–2020

Year	1990	2000	2001	2002	2003	2004	2005	2006	2007	2008	2009
E_{OC}	100	144.63	151.61	165.39	195.03	229.57	262.42	287.34	311.54	317.50	331.95
E_{AO}		112.57	120.42	127.39	143.58	160.35	175.11	189.11	210.02	229.21	265.05
Year	**2010**	**2011**	**2012**	**2013**	**2014**	**2015**	**2016**	**2017**	**2018**	**2019**	**2020**
E_{OC}	350.53	379.85	387.47	397.94	399.91	399.33	399.30	407.36	415.38	422.17	425.99
E_{AO}	275.05	306.83	318.59	335.19	338.54	343.77	347.65	357.06	368.81	377.93	380.31

Source: own study based on BP report, CO_2 EMISSIONS FROM FUEL COMBUSTION Highlights and The Observatory of Economic Complexity.

The E_{AO} was calculated by using formula (2). In the case of China, the changes in the percentage of E_{AO} CO_2 emission in 1990 were significant (table 6). In all survey years, China had higher E_{AO} than the emission in 1990. China's E_{OC} in 2000–2020 was never lower than its value in 1990. The lowest increase of E_{AO} relative to E_{OC} was in 2000 (44.63%), and the highest was 325.99% (2020). China did not fulfil the EU energy policy (reduction of CO_2 emission) in any of the survey years. The E_{AO} in

China in each survey year was lower than the their Official-Close Emission of CO_2. The smallest spread between the two of them was in the year 2000, and it was 32.06% of China's emission of CO_2 from the year 1990. The highest spread between E_{OC} and E_{AO} in China was in 2007, 101.52%. The highest E_{AO} in China was in 2020, which was 380.31% of CO_2 from 1990. This data shows that China was a substantial net exporter of CO_2 emissions in all survey years.

6. Discussion

It is worth stressing that the EU's actions in implementing the EU energy policy are limited only to the EU area. Between 2000 and 2020, China, in all survey years, had a positive trade balance with African, Asian, European, EU, and North American countries. With Oceanian countries, China had a positive trade balance from 2000 to 2008, and South American countries had a positive trade balance from 2000 to 2001. Throughout the review, the total China foreign trade balance considering five continents was negative in Oceania – 2009–2020 and South America – 2002–2020. The total China balance by 21 years of the survey was 13533 B USD. By those 21 years, the total positive China trade balance was Asia (4478 B of USD), Europe (3716 B of USD), North America (5560 B of USD), and Africa (294 B of USD). The only continents with which China had a negative trade balance from 2000 to 2020 were Oceania, which was -390 B of USD, and South America's -126 B of USD. China's trade balance had a tremendous impact on E_{AO} in all of the 78 countries that participated in the research.

What does influence on value of E_{AO}? It is Official-Close Emission of CO_2 of a particular country, the % of GDP exported goods from the China. It must be considered also the % of GDP of countries from which the China imported goods. This directly results in China's enormous influence on CO_2 emissions in other countries.

EU is the world leader in CO_2 reduction. One of its tolls is the EU energy policy. What can/should the EU do to make other countries do more to reduce CO_2 emissions? What are the challenges that the EU faces? There are two possibilities. One is doing nothing, living it without changes. The EU will be content with its energy policy, with a reduction of CO_2 emissions inside the EU. However, it will change nothing. The EU will still import CO_2 from outside the EU through products and services and continue contributing to CO_2 emissions outside the EU. The second option is to change its approach to the energy policy to be more global. The EU should take into account the CO_2 emissions that are imported into the EU. The EU should consider some instruments that encourage countries outside the EU to do similar activities to reduce CO_2 emissions. For example, ecological taxes (Fortuński 2012–2023; Bogrocz 2008; Graczyk, Jakubczyk 2005; Kaczmarski 2010; Kryk 2012a, 2012b). The E_{AO} also indicates the ineffectiveness of international agreements in reducing emissions of CO_2, such as the Kyoto Agreement.

The main challenges in the case of the second solution are retaliation activity undertaken by the countries from which the EU imports and on each of these "ecological taxes" or other instruments would be imposed. This will, among others, include transaction costs. The other challenge will be how to promote the reduction of CO_2 emissions in countries outside the EU. Because China, as we saw in previous data, imports from European countries are much smaller than China's exports to those countries, it will be easier for the EU to decrease China's CO_2 emissions. This is because China is more interested in not losing the EU as a destination for export – it would be more costly for China than for the EU to establish a new form of CO_2 tax.

The appropriate would be tariffs on all kinds of products and services imported from China to the EU. Such a tariff from the EU would likely trigger a counteraction from China in the form of tariffs on products from the EU.

Another problem is determining the reference period to which the volume of CO_2 emissions in China should be referred. Setting this to 1990 seems unrealistic. In this context, it should be noted that China's official CO_2 emissions increased throughout the period under review. The same applies to the reduction in CO_2 emissions. In the fit for 55 documents, there is talk of a 55% reduction in CO_2 emissions by 2030 (compared to 1990) and achieving climate neutrality by 2050.

About China, there is no mention of any reduction in emissions, whether compared to 1990, 2000, or even 2010. it also seems that not addressing the issue of the export and import of CO_2 emissions and its limitation by the EU on non-EU countries is deceiving EU citizens. The EU spends vast amounts of money on climate transformation. People bear the costs associated with this daily in the form of higher electricity bills and loss of jobs in emission sectors. Other countries outside the EU do not bear such expenses, which is unfair.

Another problem is that the EU is responsible for a small percentage of global CO_2 emissions – about 8%. This means that if the EU does not engage other countries in similar actions, particularly the largest CO_2 emitters, which are also its largest trading partners, climate protection actions will not bring benefits at the global level.

7 Summary

The European Union is regarded as the leader in the fight against global warming, a battle for clean energy, and a reduction of CO_2 emissions. Unfortunately, its actions are isolated, which leads to the situation that even such a large economy as the EU, which is strongly economically related to other countries through trade, can only change a little within this issue. The EU and China are leaders in world trade. The EU could use its position in international trade to achieve its own energy policy goals – reducing CO_2 emissions.

Trade relates to a balance of CO_2 hidden in goods imported to and exported from China. It affects the E_{AO} in all 78 countries from this survey. The impact of China's international trade was very high, and for most of the years, continents negatively impacted other countries. This means that China mainly exported CO_2 to those countries. It was also the case for UE countries. The impact of the trade on CO_2 emission was huge because of the substantial international trade in goods and services between China and the EU.

The effectiveness of its members implementing the EU energy policy is limited only to the EU's territory. It can result in the EU energy policy not being regarded as a SD policy and being related to high costs. If the EU is interested in reducing CO_2 emissions, it must consider it. Because of those high costs, EU countries try to reduce them by importing parts, components, and products from cheaper countries, which very often have higher emissions of CO_2. It is usually because, in those countries, environmental law is more relaxed than in the EU.

This situation brings some challenges ahead of the EU, especially in their energy policy in the CO_2 aspect. The survey has shown that the EU should consider changes in its energy policy. This policy should take into account an element of CO_2 emission more globally. The results of the research indicate that CO_2 emission is a global problem. It requires the EU to consider introducing a new instrument that would incentivize countries outside the EU to take effective action to reduce CO_2 emissions. A new instrument, an eco-energy tax, could be introduced for that purpose. The EU would apply it to all trade partners, individual countries, or groups of countries. Additionally, this tax would concern the volume of particular countries exporting to the EU (Bielecki et al. 2016: 43–46). Fortunately, EU will have some new carbon tax shortly (Salzman 2023; Carbon Border Adjustment Mechanism). Unfortunately, this kind of action could bring contractions.

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