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RESEARCH ARTICLE

ARE INFORMATION AND COMMUNICATION TECHNOLOGIES EFFECTIVE IN SERVICE EXPORTS IN SOUTH CAUCASIAN COUNTRIES?

Eser ÇAPIK*

Abstract

The advancement of information and communication technologies (ICT) in Caucasian countries has begun to have an impact on the service sector. ICT has made many firms in the service sector more competitive and innovative, enabling them to achieve high productivity at a low cost. These effects of ICT in the service sector can lead to economic growth and development in Caucasian countries.
This study examines the impact of information and communication technologies on service exports in Caucasian countries. The time period covering 2003-2022 is analyzed using time series. Given the differences in internet infrastructure, internet access, and internet availability across each country, a distinct ICT indicator is utilized for each country. The empirical findings reveal that fixed telephony contributes to service exports in Azerbaijan, service exports are influenced by individual internet usage rates in Georgia, and mobile telephony plays a role in service exports in Armenia.
Keywords: Service Export, ICT, Time Series, South Caucasian Countries
JEL codes: C3, L8, L86, N75

1. Introduction

Today, unlike many physical products, service sectors based on direct human interaction such as education, law, tourism, health and beauty services have become one of the most important sectors contributing to the development levels of countries. Rapid developments in Information and Communication Technologies in the last two decades have contributed to the fast growth of new service sectors and have significantly affected the consumption of all goods and services. In this process, it is inevitable to say that the rapidly developing internet actually plays an important role in the services traded (OECD, 2000).

In terms of information and communication technologies, especially the Internet has contributed to the globalisation of the world economy by playing an effective role in accessing information, ideas, various expertise and innovations across borders (Choi, 2010). As a result of the revolutionary developments in the ICT sector in the 90s, there have been significant changes

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in economic activities in the world such as increased diversity, productivity and global trade (Fink et al. 2005). Developments in ICT have significantly reduced the importance of physical distance in trade (Freund and Weinhold, 2004). Because the innovations in ICT have brought communication tools such as telephone, e-mail and virtual conference to the forefront and thus distance has become unimportant. In other words, Bloom, Sadun and Van Reenen (2012) argue that information and communication technologies have enabled businesses to become independent from geographical constraints by enabling remote working and virtual business environments. Thus, thanks to ICT, international communication costs are significantly reduced and international co-operation and trade are increasing (Clarke, 2008).

In fact, ICT has several mechanisms that affect trade (Nath & Liu, 2017). With these mechanisms, it facilitates both the flow of information, the automation of business processes and the increase in productivity with the effect of digitalisation, and the sharing of information and cooperation of businesses and individuals on a global scale (Jungmittag and Welfens, 2009). With advances in information and communication technologies, delays in obtaining and transmitting information can be reduced (Nath and Liu, 2017). With the development of ICT, the operational efficiency of enterprises increases and costs decrease (Brynjolfsson and Hitt, 2000), and competitive advantage can be gained by improving the decision-making processes of enterprises with big data and analytics (McAfee et al., 2012). Of course, it should be emphasised that the Internet, which has the potential to create large global markets for certain goods (Freund and Weinhold, 2002), requires a strong ICT infrastructure to function effectively. Because ICT infrastructure constitutes the basic condition for the Internet to become an effective tool for collecting, processing and disseminating information (Vemuri and Siddiqi, 2009).

With information and communication technologies (ICT), the trade in services can be revolutionised in a variety of ways. For example, ICT enables firms to reach a global customer base and facilitates services to reach potential customers around the world through the Internet, social media and digital marketing tools such as Amazon, Alibaba and Upwork. This facilitates cross-border trade in services and increases the international mobility of labour and services.

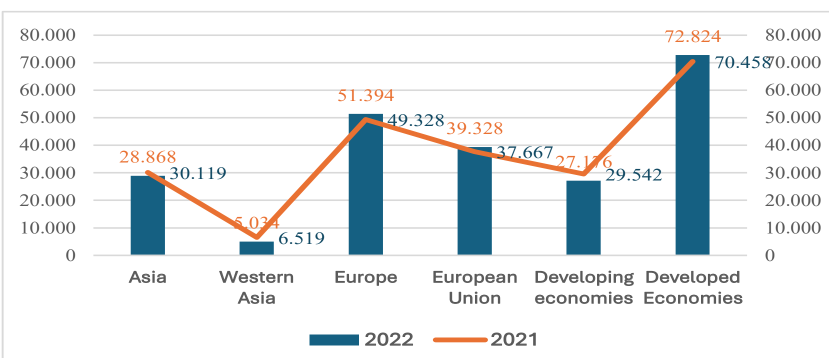


Figure 1. World Services Exports (2018-2022, %)

Source:(*UNCTADstat*, 2023), International Trade Statistics

When Figure 1 is analyzed, it can be observed that Asian countries accounted for approximately 29% of the world's service exports in 2021-2022, while West Asian countries accounted for 5%. European countries contributed to 51% of the world's service exports, with European Union countries comprising 39%. Looking at developed economies, it is evident that developing countries accounted for 73% of the world's service exports, while developed country economies accounted for 27%. In 2022, there was a noticeable increase in Asian countries and developing country economies, whereas there was a decreasing trend in the economies of other countries (UNCTAD, 2023). Of particular importance to us are the South Caucasian countries within the Western Asian country group.

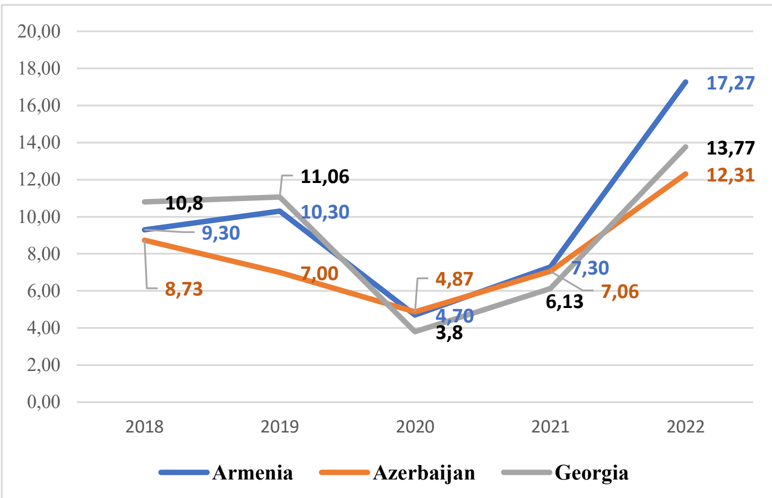


Figure 2. Exports of Services in the South Caucasus Countries (2018-2022, %)

Source: (*UNCTADstat*, 2023), International Trade Statistics

In recent years, changes in the global economy, especially driven by developments in Information and Communication Technologies (ICT), have led to economic diversification in the South Caucasus countries, which are situated in the heart of the oil-rich region, and this has had significant implications for the service sector. According to Figure 2, Georgia ranked first in service exports in 2018 with 11%, followed by Armenia in second place with 9%, and Azerbaijan in third place with 8%.

These service exports consist of various categories. The proportion of the sectors that constitute service exports in South Caucasian countries is shown in Table 1.

Table 1 Sectoral Services Exports in South Caucasus Countries (2022, $ billion)

Types of services	Azerbaijan	Armenia	Georgia
Transport services[1]	17,54491779	63,14048855	32,29859419
Travel services	45,19242181	8,24016753	48,88187694
Air transport, freight[2]	1,347896	3173,72377	0,11335
Air transport, passengers carried	179200	1102455	84591
Air transport, registered carrier departures worldwide	1598	15844	1000
Computer, communications and other services[3]	33,00553918	27,22953872	14,66987281

NOT: (% of service exports, BoP)
[2](Million ton-km)
[3] (% of commercial service exports)

Source: (*UNCTADstat*, 2023), International Trade Statistics

Table 1 shows that in 2022, transport services ranked first in Armenia's service exports, while Georgia ranked first in the travel sector and Azerbaijan ranked second. Azerbaijan ranks first in computer, communication and other services.

It can be argued that the service sector, which has recently been identified as the engine of economic growth, has demonstrated significant performance with the process of digitalization, based on innovation activities (OECD & Eurostat, 2018, p. 81). While innovation occurs through the utilization of knowledge (Antonelli, 2000), both innovation and access to knowledge are facilitated by the Internet, which serves as a powerful tool in the national economy (see, for example, Myovella et al., 2020). Data from The World Bank (2023) regarding Internet usage, which is crucial for the economy, indicates that Caucasian countries exhibit low performance in terms of Internet usage. According to the latest data published by The World Bank for 2021, it is evident that the percentage of GDP attributed to individual Internet use in Armenia, Azerbaijan, and Georgia is 0.19%, 0.18%, and 0.28%, respectively. It is believed that the level of Internet usage plays an influential role in the service sector, as it does in many other sectors.

In light of this information, to the best of our knowledge, no studies have yet addressed the impact of the ICT sector on trade in services in the South Caucasus. Therefore, this study aims to analyze the impact of ICT, which is at a developing level in South Caucasus countries (Doyar et al., 2023), on the success of service exports. These exports constitute an important pillar of economic growth in South Caucasus countries (Çapık & Ören, 2023).

Detailed answers to the question of the determinants of Internet use in the region are discussed in (Doyar et al., 2023). In this study, total service sector exports in South Caucasian countries, which have not been examined before, will be analysed with the effect of ICT determinants in this study and the weakness of the literature will be strengthened.

The paper is organised as follows: in the first part, the history of the service sector and ICT is briefly reviewed and the situation of the South Caucasus countries is discussed. The second section provides a literature review on the subject. After the empirical in the third section, the data, model and methodology are explained in the fourth section. Following comparing the findings with the literature in the fifth section, the last section discusses the policy implications of the findings and presents various conclusions.

2. A Brief Historical Overview of the Service Sector and Information and Communication Technologies

The development process of the economy in history includes a very interesting story. Trade in services has played an important role in this story of the economy and the way of thinking on the subject has changed over time. Looking at the economic history of the countries, after the first and second industrial revolutions, there has been an important shift from agriculture to industry, and in the last fifty years, a shift from industry to services has been realised. Fisher (1939) and Clark (1941), two prominent economists of the 1930s, developed models for the stages of economic growth. In these models, the production of raw materials is defined as "primary", manufacturing goods as "secondary" and the production of services, which play an important role in the development of a country, as "tertiary" (Grubel & Walker, 1989; Kuźnar, 2016; Shanmugam & Latha, 2014).

Since the realisation of any goods production depends on raw materials, raw material production is defined as the primary sector, the manufacturing sector is defined as the secondary sector since it also depends on raw materials, while the production of services is defined as the tertiary sector since it depends on both the primary and secondary sectors (Shanmugam and Latha, 2014). Based on differences in product labour productivity and the size of the labour force in various economic activities, (Clark, 1941, p. 121), defined the primary sector as activities that use and transform natural resources (agriculture, forestry and fishing), the secondary sector as production activities that continuously transform natural resources into transportable products, and the tertiary sector as service activities (consumer and producer services, construction and labour productivity per worker). Important economists such as Adam Smith, David Ricardo and Karl Marx, while studying labour productivity and valorisation, treated services differently from goods and products.

In the period of the industrial revolution, the labour force in the agricultural sector shifted towards the manufacturing sector and it was observed that human needs reached a saturation point, incomes increased and the demand for food and goods consumption decreased. With the increase in consumer income, there has been an orientation towards preferences for leisure and entertainment (Hospers, 2004, p. 12). Consequently, there has been an evolution towards the service sector. As such, the growth of the service sector in most developed countries has been associated with the level of income (Ramakrishna, 2010; Schettkat, 2007; Witt & Gross, 2020). The service sector, whose contribution to economies has steadily increased over time, contributes to productivity and economy-wide growth as it provides basic inputs to other products and services (Çapık and Ören, 2023). The service sector, which is now recognised as the fastest growing area of international trade (Bradley et al., 1995), as (Vandermerwe & Chadwick, 1989, p.

79) put it: “the whole world is today a field of service activities”. However, technological advances and globalisation in the last decade have radically changed the definition of services and the way economists view services.

In the service sector, which has become a major economic centre, most of the economic value produced is based on information. Due to significant advances in information and communication technology (ICT), it has recently come to be recognised as an innovative sector and has been the subject of considerable trade (Miles, 2000). ICT has changed the nature of the production of services as well as the nature of service exports in particular, paving the way for both a rapid increase in service exports and a significant share of services in the GDP pie.

The history of ICT dates back to the time when people started to use objects to communicate with each other. In other words, ICT started with the rise of mankind (Duque et al., 2006, p. 33-39). It can be said that the basic concept of ICT, which appeared clearly in the 1970s, actually dates back to the Second World War alliance of the military and industry, which played an active role in the development of electronics, computers and information theory. Going back a little further, after the 1940s, the military was the main source of R&D funding for the expansion of automation in order to replace manpower with machines. After the 1950s, computer types started to develop, which became more capable but smaller in size. The study, design, development, support or management of computer hardware constitutes information and communication technologies (ICT) (Sakenov, 2018). Information and communication technologies include computers, the Internet, broadcasting technologies (radio and television) and telephones (UNDP, 2001, p. 29) As a result of the closeness of information and communication, a technological revolution has been created, and the most striking aspect of this situation is that it has revealed important results for its application in different fields of economic activities.

Information and communication technologies have revolutionised the way most traditional services are produced and sold, offering various opportunities in different areas of the service industry and playing an important role in firms’ innovation activities (Evangelista & Sirilli, 1995). Services now include digital and remote services such as online shopping, digital education, telecommunications, consultancy, software development and many many other different areas. These technological changes have contributed to the transformation of services into a sector that is not only based on physical interactions, but also has a major impact on global markets and has become a fundamental component of the economy (Baumol, 1967). Information technologies are in the service sector, especially in the application of sub-sectors such as financial services and communication services (Biswas, 2020). It can be said that information and communication technology has also made itself more and more felt in the internationalisation of services.

In recent years, the changes that have occurred with the global economy, especially in line with the developments in Information and Communication Technologies (ICT), have started to show themselves in the South Caucasus Countries, which are located in the centre of oil.

2.1 Armenia

ICT in Armenia, which is one of the leading information technology countries in the Commonwealth of Independent States and the Middle East, suffered a major blow with the collapse of the Soviet Union but managed to overcome this by going through a recovery process in the mid-1990s. In this case, especially in the 1990s, the opening of branches of US software companies in Yerevan, the capital of Armenia, played an active role. Thus, a new era began in Armenia with the US diaspora (Vardanyan & Sarkisyan, 2004). The ICT industry globally. The ICT industry of Armenia is demonstrating its potential in different service sectors, drawing the interest of investors, policy makers and professionals (USAID, 2009, p. 12-13).

Armenia, formerly described as the "Silicon Valley of the Soviet Union", received backing from organisations including the United States Agency for International Development (USAID) and the World Bank (Krikorian, 2010). In recent years, the information and communications technology (ICT) industry in Armenia has experienced significant growth and has emerged as a highly dynamic sector of the economy (Vardanyan & Sarkisyan, 2004, p. 7-8)

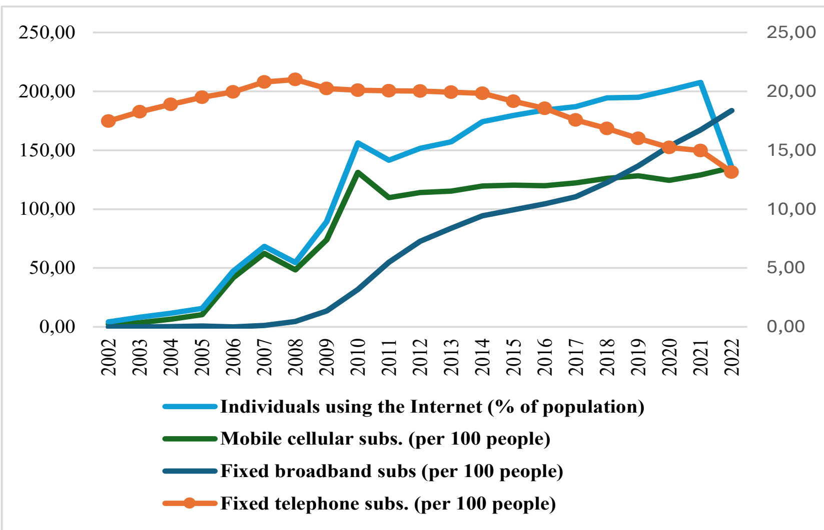


Figure 3. Timeline of the ICT Indicators in Armenia (%)

Source: Visualised using data from (ITU, 2023a, 2023b, 2023c, 2023d)

Figure 3 reveals that the proportion of internet users in Armenia's population was 5% in 2005 and soared to 55% in 2015 and to 79% in 2021. Mobile networks accounted for 110% of the total internet access in 2011, and this percentage increased to 120% in 2015, reaching 129% by 2021 (ITU, 2018). Fixed broadband subscriptions were responsible for only 0.03% of internet access in 2005, reaching 8.38% in 2014, and this rate further surged to 17% in 2021 (ITU, 2023). Finally, while 17% of internet access was provided by fixed telephone subscriptions in 2002, this rate decreased to 20% in 2012 and to 13% in 2022. In fact, it is possible to say that the demand for mobile phones has an effect on this situation

2.2. Georgia

In Georgia, computer technologies were introduced in the late 1960s when the country was under Soviet Union rule. Numerous computer centres were established in ministries, factories, universities, and research institutions during this period. The economic crisis triggered by the Soviet Union's collapse led to swift closures of computer centres. In 1995, with backing from the United States, the American NGO Parliamentary Human Rights Foundation facilitated the first Georgian Parliament Internet access. Consequently, the official parliamentary website became the inaugural site (Karumidze, 2001). With the escalating use of the Internet and personal computers in Georgia, the 1980s saw the official introduction of personal computers that were actively employed in accounting services. The development of the Internet in Georgia is closely linked to both local content websites and Internet access (Karumidze, 2001).

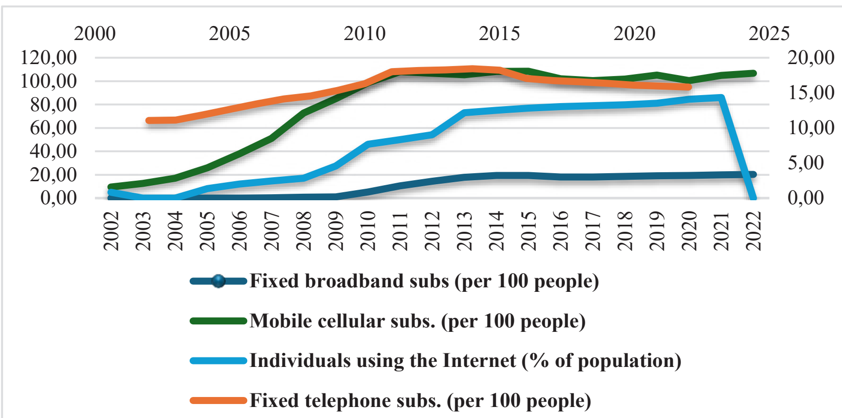


Figure 4. Timeline of the ICT Indicators in Georgia (%)

Source: Visualised using data from (ITU, 2023a, 2023b, 2023c, 2023d)

An of Figure 4 shows that in 2005, around 4% of the population in Georgia had access to the internet. This percentage increased to 20% in 2010 and further to 76% in 2021. In 2000, mobile networks provided 5% of internet access, rising to 110% in 2010 and 148% in 2021 (ITU, 2023). Fixed broadband subscriptions accounted for approximately 5% of internet access in 2010, rising to 27% by 2021 (ITU, 2023). In 2002, 16% of internet access was provided by fixed telephone subscription, while this rate dropped to 8% in 2022.

2.3. Azerbaijan

Information and communication technologies (ICT) are considered integral to the development of the Republic of Azerbaijan. The National Strategy highlights its significant role. Since 1991, when internet technology was introduced in Azerbaijan, international connectivity was established in 1993 with the help of British Petroleum and Turkey. The first Azerbaijani website was launched in 1994 (MINCOM, 2021).

Permanent internet access was provided to Azerbaijan through the Azerbaijani Academy of Sciences in 1995. Despite undergoing challenging processes in information and communication technologies, Azerbaijan's most unfavorable aspect in terms of internet accessibility is the unsuitable state of the country's telecommunications infrastructure. Although UNDP, Soros Foundation, IREX and NATO have endeavored to assist in the development of the country's internet infrastructure, positive results have not been achieved. When compared to the 1990s, it can be said that Azerbaijan has made significant progress in terms of internet development (Izmaylov, 2001). In the mid-2000s, there was a rapid increase in mobile cellular subscriptions. In the years 2005-2008, the State Program on the development of information and communication technologies was adopted in the Azerbaijan Republic. Following the launch of 3G services in 2009, there was an expansion of mobile broadband subscriptions (MINCOM, 2021).

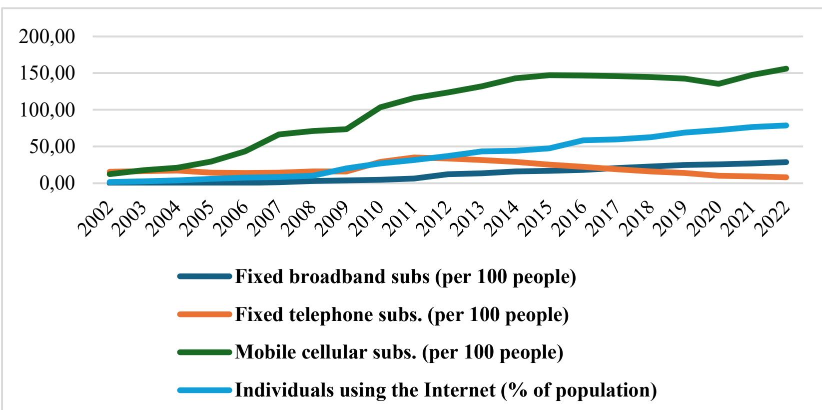


Figure 5. Timeline of the ICT Indicators in Azerbaijan (%)

Source: Visualised using data from (ITU, 2023a, 2023b, 2023c, 2023d)

Looking at Figure 5, we can see that in 2005, 8% of Azerbaijan's population were internet users. This figure increased to 50% in 2011 and will reach 86% in 2021. In 2010, 5% of internet access in Azerbaijan was provided through fixed broadband subscriptions, while this figure has risen to 20% towards 2021. Also, it can be observed that the proportion of internet usage through mobile networks has grown from 26% in 2005 to 105% in 2021 (ITU, 2023). In 2002, 11 per cent of internet access was provided by fixed telephone subscription, while this rate increased to 16 per cent in 2022.

3. Literature Review

With the development of technology, empirical studies on the determination of business performance (Wakelin, 1998) generally stem from two different theoretical traditions. The first is based on the Heckscher-Ohlin-Samuelson trade model in which trade is determined by the relative factor. The second theoretical tradition is the Technology Gap Theory of Trade (Posner, 1961) and the Product Cycle Approach to Trade (Vernon, 1966). According to Kaldor's empirical evidence, sometimes price-related factors are insufficient for exports. Therefore, differences in the price level are not sufficient to explain the differences in export performance between countries. In this case, it is stated that non-price factors should be considered (Visús & Zayas, 2003).

Onyeiwu (2002) analysed the extent to which technology, one of the non-price factors, is integrated into the global economy of a country. Freund and Weinhold (2002) argue that the Internet, the most important determinant of technology, can help create global markets for traded goods by reducing fixed costs. Chatti & Haitham Khoj (2020) aimed to examine the causal relationships between service exports and internet penetration for 116 countries in the period 2000-2017, and concluded that there is a bi-directional causality between service exports and internet adoption for developed countries. Nasir and Kalirajan (2016) examined the export performance of developed Asian economies in selected modern services and found that the ICT and innovation performance of the emerging economies in South Asia and the Association of Southeast Asian Nations in realising the export potential of the emerging economies in South Asia and the Association of Southeast Asian Nations is quite weak compared to the developed economies in North America and Europe. It has shown that ICT infrastructure quality is among the key factors in realising service export potential.

Aydınbaş and Erdinç (2023) investigated the effects of information and communication technologies on human capital and found that the share of ICT exports in total exports increases human capital. Kırca and Akkuş (2020) examined the effects of internet use and economic growth on electricity consumption in EU-15 countries, and the results showed that for all EU-15 countries, the change in internet use can decrease electricity consumption very slightly, but the change in economic growth can increase electricity consumption. Ozcan (2018) analysed the effects of information and communication technologies (ICT) on international trade between Turkey and its trading partners and the results showed that ICT has positive and significant effects on both import and export volumes of Turkey.

Table 2. Literature Review

Author	Country	Method	Period	Variables												
				lngdp	lngdpp.	enxhange	Pop	lnCO emsyn	unemp	Emp.	Lang.	lndist.	ICT (1)	(2)	(3)	(4)
Nath & Liu (2017)	49 developed and developing countries	Panel data	2000-2003	-	(+)***	-	(-)*	-	-	-	-	-	(+)	(-)	(+)**	(+)
Wang & Choi (2019)	BRICS country	Panel data	2000-2016	(+)***	-	-	(+)	-	(+)***	-	-	-	(+)***	(+)	(+)***	(+)***
Lichy et al. (2022)	44 high-income 36 low – and middle-income countries	Panel data	2000-2019	(+)***	-	(+)***	-	-	-	-	-	-	-	-	(+)***	(+)***
Aijaz et al. (2022)	China	Zaman serisi	1990-2020	-	-	-	-	(+)***	-	(+)	-	-	(-)	-	(-)***	-
Nasir & Kalirajan (2016)	Asean countries	Panel data	2005-2017	(+)***	-	-	-	-	-	-	(+)***	(-)***	-	(+)***	(+)***	-
Wardani et al. (2020)	Asean countries and Indonesia	Gravity model	2005-2017	(+)***	-	-	-	-	-	-	(+)***	(-)	-	(+)***	-	-
Freund & Weinhold (2002)	The US	Panel data analysis	1995-1999	(+)**	-	(+)	(-)***	-	-	-	-	(-)	-	(+)***	-	-
Choi (2010)	151 countries	Panel data analysis	1990-2006	(+)***	-	-	(-)***	-	-	-	-	-	-	(+)***	-	-
Mattes et al. (2012)	The US	Panel data analysis	1995-2007	(+)***	-	-	-	-	-	-	(+)	(-)***	-	-	-	-
Clarke & Wallasten (2006)	Developing countries, developed countries, and total exports (i.e., to all countries)	OLS and 2SLS	2001	(-)***	-	-	(+)*	-	-	-	-	(-)***	-	-	(+)***	-
Clarke (2002)	Low and middle-income economies in Eastern Europe	OLS	1999	-	(-)	-	(-)	-	-	-	-	(+)	-	(+)***	-	-

Vemuri & Siddiqi, (2009) in the study examining the impact of information and communication technology (ICT) and the Internet on international trade for 64 countries from 1985 to 2005, it was found that the ICT infrastructure and the Internet's openness to commercial transactions had a positive and significant effect on the international trade volume. In Clarke and Wallasten (2006), it was concluded that access to the Internet improves export performance in developing countries. The study showed that when Internet penetration is high in developing countries, more exports are made to developed countries. In the empirical literature, ICT infrastructure, particularly the Internet, has been found to have a significant impact on trade in services (Choi, 2010). Liu and Nath (2017) examined the relationship between the Internet and trade in services in developing countries over the period 1995-2010, and the results showed that Internet penetration positively affects trade in developing countries. Luong and Nguyen (2021) investigated the effects of information and communication technology (ICT) on trade in services and found that this effect yields similar results on both imports and exports.

Although there are many factors affecting service exports such as trade agreements, infrastructure development, human capital, etc., the model was constructed based on the variables used in the literature studies in Table 2. Table 2 summarises the relevant literature by showing the separate effects of each of the variables usually targeted for trade in services, such as GDP, GDP per capita, population, international distance, technological skills and technological ownership.

4. Research Design

This study aims to contribute to the literature by utilizing the most up-to-date datasets. Firstly, it considers a comprehensive dimension of ICT for South Caucasus countries, taking into account their specific characteristics. What sets our study apart from others is its approach, which examines service exports in three different countries using three separate models and various ICT indicators tailored to their specific characteristics.

This paper analyses the role of the ICT index in services exports for the economies of Armenia, Georgia and Azerbaijan for nineteen consecutive years from 2003 to 2022. The selected years cover global trends such as the rise of the Internet. The choice of period is based on data availability, and the data used in this study were obtained from the World Bank's World Development Indicators (WDI) database (2023). Data collection methods may differ in some countries, and missing or outdated data may lead to incorrect relationships in analyses of services exports. However, reliable and consistent data on ICT and services exports between 2003 and 2022 for all three countries are available at the World Bank. The data set is based on time series and the variables used in the study are presented in Table 3.

Table 3. Variable Definitions

Variables	Variable definitions	Data
Log of SE	Service exports (BoP, current USD)	WDI (2023)
Log of GDP	Per capita GDP (constant, 2010 USD)	WDI (2023)
Log of POP	Population (total)	WDI (2023)
Log of Offical Exchange rate	1 US$ = country currency	WDI (2023)
Log of individuals using the internet	(% of population)	WDI (2023)
Log of fixed-telephone subscriptions	(per 100 inhabitants)	WDI (2023)
Log of mobilecellular telephone subscriptions	(per 100 inhabitants)	WDI (2023)

Kaynak: World Bank, WDI (2023)

4.1. Model and Methodology

Following Clarke and Wallasten's (2005) of ICT indicators as determinants of service export, we can present the following model.

$$SE_t = \alpha_0 + \beta 1 lngdp_t + \beta 2 lnpop_t + \beta 4 exc_t + \beta 5 ict_t + \varepsilon_t$$

SE, which is included in the model as the dependent variable, is considered as total service exports. Subsequently, the control variable lngdp represents the per capita GDP of countries at time *t* in constant prices. Individuals who are socio-economically privileged, meaning financially well-off, have a greater likelihood of adopting and benefiting from technological innovations such as the internet (Mills & Whitacre, 2003). Although income facilitates Internet access, it cannot be considered as the sole driving force behind Internet usage. Hence, other variables have been included in the model by drawing inspiration from the literature. Specifically, lnpop has been incorporated as the final variable, reflecting the total population of countries at time *t*. As population and density increase, internet usage also increases, leading to lower internet access costs (Stoneman & Karshenas, 1995). exc also represents the official exchange rate of a country at a given point in time. Since the exchange rate is a determining factor for international trade activities, the relationship between international trade and exchange rate has been extensively researched by scholars (Lichy et al., 2022). The model includes ICT as the dependent variable, which comprises fixed-line subscriptions (per 100 inhabitants), mobile phone subscriptions (per 100 inhabitants), and the percentage of individuals using the Internet. When referring to service exports between 2003-2022, t represents time, and ε indicates the error term.

To identify the stationary states of the series, we utilized the Augmented Dickey-Fuller (ADF) unit root test (Dickey & Fuller, 1981) in our study. Because non-stationary time series can lead to the problem of spurious regression (Granger & Newbold, 1974, p. 111), this can risk showing a relationship between variables that do not exist and can also result in the misinterpretation of the coefficients. Therefore, we performed the Augmented Dickey-Fuller (ADF) test as a first step to determine the stationarity of the variables. The functional form of ADF can be expressed as follows (Malik & Velan, 2020):

$$\Delta Y_t = \alpha + \emptyset Y_{t-1} + \sum_{i=1}^{m} \beta_i \Delta Y_{i-t} + et \qquad (1)$$

$H_0 = \beta_i = 0$ (*i.e., time series has unit root)*

$H_1 = \beta_i \neq 0$ (*i.e., time series has not unit root)*

In order to carry out the causal in the next step, we require the value of $\boldsymbol{m_{max}}$, which represents the maximum integrated order obtained from the unit root test.

It is necessary to establish their co-integration in order to gain insight into the causal relationship between service exports and ICT indices. Since the model includes more than two variables, the Johansen cointegration technique developed by Johansen (1988) and Johansen and Juselius (1990) is used. When more than two variables are involved, more than one long-term equilibrium relationship may emerge. As a first step in the cointegration test, all variables included in the model must not be stationary at the level, but rather become stationary when first differenced. This model is not applicable when there are varying levels of stationarity. To apply this model, a VAR model must be established first, and the lag order should be determined using Akaike and Schwarz information criteria. Once the model with the lowest AIC and SBC was selected, it was identified as the optimal lag length and denoted as p.

In our study, we utilize the Toda-Yamamoto causality test (TY) developed by Toda and Yamamoto (1995). In Granger causality analysis, if variable X yields better results than predicting variable Y using all available information, it suggests causality from X to Y (Granger, 1969). Additionally, Toda and Yamamoto (1995) describe the estimation of VAR models for integrated or cointegrated series at various levels and testing parameter matrix restrictions. For the purpose of conducting the test, the maximum integration degree $\boldsymbol{m_{max}}$of the variables has been determined using unit root tests. Information criteria were used to determine the optimal lag length, denoted as p, for the VAR model, a multivariate time series model proposed by Sims (1980). Later, a VAR model with $\mathbf{p}^{+}$ lags was forecasted. The direction of causality was determined by applying a Wald test with an asymptotic chi-square distribution to the p-lag values to determine whether the coefficients were statistically different from zero.

When the $\boldsymbol{p}$ optimum delay length, $\boldsymbol{m_{max}}$ maximum integration order, and $\boldsymbol{u}$ white noise term are used, the VAR model to be estimated for the *lngdp, lnpop, lnict1, lnict2, lnict3,* and *lnexc* variables for the SE procedure can be written as follows:

$$
\begin{aligned}
loggdp_t = \alpha_0 &+ \sum_{i=1}^{p} \beta_{1i}\, logpop_{t-i} + \sum_{j=p+1}^{m_{max}} \beta_{2j}\, logpop_{t-j} \\
&+ \sum_{i=1}^{p} \delta_{1i} exc_{t-i} + \sum_{j=p+1}^{m_{max}} \delta_{2j}\, exc_{t-j} \\
&+ \sum_{i=1}^{p} x_{1i} ict1_{t-i} + \sum_{j=p+1}^{m_{max}} x_{2j}\, ict1_{t-j} \\
&+ \sum_{i=1}^{p} \vartheta_{1i} ict2_{t-i} + \sum_{j=p+1}^{m_{max}} \vartheta_{2j}\, ict2_{t-j} \\
&+ \sum_{j=p+1}^{m_{max}} \partial_{1i} ict3_{t-i} + \sum_{j=p+1}^{m_{max}} \partial_{2j} ict3_{t-j} \\
&+ u_{1t}
\end{aligned}
\qquad ((2)
$$

$$
\begin{aligned}
logpop_t = \alpha_0 &+ \sum_{i=1}^{p} \rho_{1i} logdp_{t-i} + \sum_{j=p+1}^{m_{max}} \rho_{2j}\, loggdp_{t-j} \\
&+ \sum_{i=1}^{p} \delta_{1i} exc_{t-i} + \sum_{j=p+1}^{m_{max}} \delta_{2j}\, exc_{t-j} \\
&+ \sum_{i=1}^{p} x_{1i} ict1_{t-i} + \sum_{j=p+1}^{m_{max}} x_{2j}\, ict1_{t-j} \\
&+ \sum_{i=1}^{p} \vartheta_{1i} ict2_{t-i} + \sum_{j=p+1}^{m_{max}} \vartheta_{2j}\, ict2_{t-j} \\
&+ \sum_{j=p+1}^{m_{max}} \partial_{1i} ict3_{t-i} + \sum_{j=p+1}^{m_{max}} \partial_{2j} ict3_{t-j} \\
&+ u_{1t}
\end{aligned}
\qquad ((3)
$$

$$
\begin{aligned}
logexc_t = \alpha_0 &+ \sum_{i=1}^{p} \rho_{1i} logdp_{t-i} + \sum_{j=p+1}^{m_{max}} \rho_{2j}\, loggdp_{t-j} \\
&+ \sum_{i=1}^{p} \delta_{1i} pop_{t-i} + \sum_{j=p+1}^{m_{max}} \delta_{2j}\, pop_{t-j} \\
&+ \sum_{i=1}^{p} x_{1i} ict1_{t-i} + \sum_{j=p+1}^{m_{max}} x_{2j}\, ict1_{t-j} \\
&+ \sum_{i=1}^{p} \vartheta_{1i} ict2_{t-i} + \sum_{j=p+1}^{m_{max}} \vartheta_{2j}\, ict2_{t-j} \\
&+ \sum_{j=p+1}^{m_{max}} \partial_{1i} ict3_{t-i} + \sum_{j=p+1}^{m_{max}} \partial_{2j} ict3_{t-j} \\
&+ u_{1t}
\end{aligned}
\qquad ((4)
$$

$$logict1_t = \alpha_0 + \sum_{i=1}^{p} \rho_{1i} logdp_{t-i} + \sum_{j=p+1}^{m_{max}} \rho_{2j}\, loggdp_{t-j} + \sum_{i=1}^{p} \delta_{1i} pop_{t-i} + \sum_{j=p+1}^{m_{max}} \delta_{2j}\, pop_{t-j} + \sum_{i=1}^{p} x_{1i} exc_{t-i} + \sum_{j=p+1}^{m_{max}} x_{2j}\, exc_{t-j} + \sum_{i=1}^{p} \vartheta_{1i} ict2_{t-i} + \sum_{j=p+1}^{m_{max}} \vartheta_{2j}\, ict2_{t-j} + \sum_{j=p+1}^{m_{max}} \partial_{1i} ict3_{t-i} + \sum_{j=p+1}^{m_{max}} \partial_{2j} ict3_{t-j} + u_{1t} \qquad ((5)$$

$$logict2_t = \alpha_0 + \sum_{i=1}^{p} \rho_{1i} logdp_{t-i} + \sum_{j=p+1}^{m_{max}} \rho_{2j}\, loggdp_{t-j} + \sum_{i=1}^{p} \delta_{1i} pop_{t-i} + \sum_{j=p+1}^{m_{max}} \delta_{2j}\, pop_{t-j} + \sum_{i=1}^{p} x_{1i} exc_{t-i} + \sum_{j=p+1}^{m_{max}} x_{2j}\, exc_{t-j} + \sum_{i=1}^{p} x_{1i} ict1_{t-i} + \sum_{j=p+1}^{m_{max}} x_{2j}\, ict1_{t-j} + \sum_{j=p+1}^{m_{max}} \partial_{1i} ict3_{t-i} + \sum_{j=p+1}^{m_{max}} \partial_{2j} ict3_{t-j} + u_{1t} \qquad ((6)$$

$$logict3_t = \alpha_0 + \sum_{i=1}^{p} \rho_{1i} logdp_{t-i} + \sum_{j=p+1}^{m_{max}} \rho_{2j}\, loggdp_{t-j} + \sum_{i=1}^{p} \delta_{1i} pop_{t-i} + \sum_{j=p+1}^{m_{max}} \delta_{2j}\, pop_{t-j} + \sum_{i=1}^{p} x_{1i} exc_{t-i} + \sum_{j=p+1}^{m_{max}} x_{2j}\, exc_{t-j} + \sum_{i=1}^{p} x_{1i} ict1_{t-i} + \sum_{j=p+1}^{m_{max}} x_{2j}\, ict1_{t-j} + \sum_{i=1}^{p} \vartheta_{1i} ict2_{t-i} + \sum_{j=p+1}^{m_{max}} \vartheta_{2j}\, ict2_{t-j} + u_{1t} \qquad ((7)$$

The inclusion of lagged values of dependent variables in the VAR model, in which all variables are lagged along with their own lags, makes it possible to make strong forecasts for the future (Kumar et al., 1995). In VAR models, which are used when there is no cointegration between variables and the series are stationary, the error correction model (VEC model) is estimated if the variables are cointegrated. One of the biggest advantages of the VAR model is that it takes into account the effect of all variables included in the model and performs the co-integration test without any restrictions.

4.2. Empirical Results

This section presents the results obtained from unit root and causality tests for each of the three countries separately.

Results for Azerbaijan

Table 4. Unit Root Test Results

ADF (Constant)				ADF (Constant and trend)		
	Level	1^{st} $dif.$	2^{nd} $dif.$	Level	1^{st} $dif.$	2^{nd} $dif.$
LogSE	-2.034	-2.709*	-5.315***	-1.421	-2.789	-5.311***
LogGDP	-3.579	-1.571	-3.787***	-1.185	-2.088	-4.523***
LogPOP	-1.924	-1.350	-4.112***	-0.796	-1.823	-5.029***
LogICT1	-1.932	-2.035	-6.181***	-0.228	-2.922	-6.022***

*** and * represent significance at 1% and 10%, respectively. Lag length for ADF test is chosen by Schwarz Information Criteria.

The unit root test results in Table 4 indicate that our series are not stationary at the same level. According to the ADF unit root test results in the table, all variables are stationary at the second level. For the equation with constant and trend, all variables become stationary at the second difference at 1% significance level. Considering all these outputs, the maximum order of integration is $\boldsymbol{m_{max}}$=2. Therefore, the number of additional lags for the VAR model to be estimated by stationarising is set as 2.

The next step is to determine the lag length of the model. The lag value determined with the help of Akaike Information Criterion (AIC), Schwart Criterion (SC), and Hannan Quinn Criterion (HQ) is shown in Table 5. According to Table 5, the fact that the maximum number of stars is two led us to determine our lag value as $\boldsymbol{m=2}$ according to SC and HQ values. Therefore, two lag length will be used in our model.

Table 5. Determination of the Appropriate Lag Length

Lag	LogL	LR	FPE	AIC	SC	HQ
0	66.12224	NA	2.36e-08	-6.212224	-6.013077	-6.173348
1	183.8120	176.5346*	9.43e-13	-16.38120	-15.38546*	-16.18682
2	207.4962	26.05265	5.53e-13*	-17.14962*	-15.35730	-16.79974*

* indicates lag order selected by the criterion
LR: sequential modified LR test statistic (each test at 5% level)
FPE: Final prediction error
AIC: Akaike information criterion
SC: Schwarz information criterion

After determining the lag length, the VAR (4) model with $\boldsymbol{p+m_{max}=4}$ lags was estimated and the Wald test with $\boldsymbol{p=2}$ lags was applied to perform the Toda-Yamamoto causality test in order to observe the causality relationship between the variables.

Table 6. Causality Test Results

H_0	χ^2	Prob.	Decision
LOGICT1 does Granger-cause LOGSE	5.429815	0.0662	logICT1→logSE
All does Granger-cause LOGSE	14.63776	0.0233	All → logSE
LOGSE doesn't Granger-cause LOGICT1	0.342912	0.8424	logSE ··· logICT1
All doesn't Granger-cause LOGICT1	5.711052	0.4563	All ··· logICT1

A → B means causality runs from A to B.
A ··· B means no causality between A and B

Table 6 shows that in the model where the dependent variable is total service exports, fixed telephone internet subscription, which is the variable used as an ICT indicator, causes service exports, and at the same time, there is a causality from all variables in the model towards service exports. In the model where the dependent variable is ICT1, it is seen that there is no causality from service exports to ICT1 and there is no causality from all variables to ICT1. In other words, it is determined that there is a unidirectional causality relationship between service exports and ICT for Azerbaijan.

Results for Georgia

Table 7. Unit Root Test Results

ADF (Constant)				ADF (Constant and trend)		
	Level	1^{st} $dif.$	2^{nd} $dif.$	Level	1^{st} $dif.$	2^{nd} $dif.$
LogSE	-2.578	-4.248***	-4.1525**	1.182	-3.874***	-3.923***
LogGDP	-1.835	-3.816***	-4.367***	-2.148	-3.875***	-3.238**
LogPOP	-1.974	-1.738	-5.423***	-0.738	-2.583	-5.175***
LogICT2	-5.262***	-2.412	-4.133***	-1.463	-4.159***	-4.119***

*** and * represent significance at 1% and 10%, respectively. Lag length for ADF test is chosen by Schwarz Information Criteria.

As can be seen from the unit root test results, the series are not stationary at the same level. Regarding the ADF unit root test results given in the table, it is observed that the unit root hypothesis of the LOGICT2 variable is accepted at the 1% significance level in the first difference for equations with a constant. LOGGDP and LOGGSE variables become stationary at the 1% significance level in the first difference, and all variables are accepted at the 1% significance level in the second difference and become stationary. For the equation with a constant and trend, it is observed that LOGSE, LOGGDP, and LOGICT2 become stationary at the 1% significance level when the first difference is applied, and all variables become stationary at the 1% significance level in the second difference.

Considering all these outputs, the maximum order of integration is $\boldsymbol{m_{max}=2}$ Therefore, the number of additional lags for the VAR model to be estimated by stationarising is set as 2. As the second step, the lag value determined with the help of Akaike Information Criterion (AIC), Schwart Criterion (SC), and Hannan Quinn Criterion (HQ) is shown in Table 8. According to the data in Table 8, since the maximum number of stars is two, our lag value is determined as two according to SC and HQ values.

Table 8. Determination of the Appropriate Lag Length

Lag	LogL	LR	FPE	AIC	SC	HQ
0	81.21459	NA	5.21e-09	-7.721459	-7.522312	-7.682583
1	170.3740	133.7392	3.62e-12	-15.03740	-14.04167	-14.84303
2	207.6558	41.00990*	5.44e-13*	-17.16558*	-15.37326*	-16.81570*

* indicates lag order selected by the criterion
LR: sequential modified LR test statistic (each test at 5% level)
FPE: Final prediction error
AIC: Akaike information criterion
SC: Schwarz information criterion

As seen in Table 8, the optimum lag length for the VAR model was determined as $\boldsymbol{p=2}$ by using a number of information criteria.

Table 9. Causality Test Results

H_0	χ^2	Prob.	Decision
LOGICT2 doesn't Granger-cause LOGSE	3.378351	0.3369	logICT2···logSE
All does Granger-cause LOGSE	15.99780	0.0669	All → logSE
LOGSE does Granger-cause LOGICT2	13.56926	0.0036	logSE → logICT2
All does Granger-cause LOGICT2	20.88649	0.0132	All → logICT2

A → B means causality runs from A to B.
A ··· B means no causality between A and B

In Table 9, in the model where the dependent variable is total service exports, it is seen that there is no causality from the number of individual internet users in the model as an ICT indicator to service exports, but there is a causality from all variables to service exports. In the model where the dependent variable is the number of internet users, it is seen that there is a causality from the service sector to ICT2 and again there is a causality from all variables to ICT2. For Georgia, it is determined that there is a unidirectional causality between the number of internet users and service exports, that is, the number of internet users does not cause service exports, but service exports cause internet usage.

Results for Armenia

Table 10. Unit Root Test Results

ADF (Constant)				ADF (Constant and trend)		
	Level	1^{st} $dif.$	2^{nd} $dif.$	Level	1^{st} $dif.$	2^{nd} $dif.$
LogSE	-4.293***	-4.104***	-4.573***	-1.461	-4.549***	-4.117***
LogGDP	-2.523	-3.205***	-5.586***	-3.002	-3.182	-5.579***
LogPOP	-3.541***	-0.796	-3.384***	-0.703	-2.569	-3.401***
LogICT3	-4.627***	-1.640***	-6.377***	-3.136	-2.065	-5.067***

*** and * represent significance at 1% and 10%, respectively. Lag length for ADF test is chosen by Schwarz Information Criteria.

As can be seen from the unit root test results, our series are not stationary at the same level. Regarding the ADF unit root test results in the table, it is seen that the variables other than LOGGDP are stationary for the equations with constant, the variables other than LogPOP are stationary in the first difference, and all variables are stationary in the second difference and accepted at 1% significance level. For the equation with constant and trend, it is observed that all series are non-stationary at level values, only LOGSE becomes stationary when the first difference is applied, and all variables become stationary at 1% significance level at the second difference. Considering all these outputs, the maximum order of integration is $\boldsymbol{m_{max}=2}$

The next step is to determine the lag values of the model. The lag values determined with the help of Akaike Information Criterion (AIC), Schwart Criterion (SC), and Hannan Quinn Criterion (HQ) are shown in Table 11. According to the data in Table 11, since the maximum number of stars is two, our lag value is determined as two according to SC and HQ values.

Table 11. Determination of the Appropriate Lag Length

Lag	LogL	LR	FPE	AIC	SC	HQ
0	52.13991	NA	9.54e-08	-4.813991	-4.614844	-4.775115
1	158.7183	159.8676	1.16e-11	-13.87183	-12.87610	-13.67745
2	184.3062	28.14666*	5.62e-12*	-14.83062*	-13.03830*	-14.48074*

* indicates lag order selected by the criterion
LR: sequential modified LR test statistic (each test at 5% level)
FPE: Final prediction error
AIC: Akaike information criterion
SC: Schwarz information criterion

As shown in Table 11, the optimum lag length for the VAR model was determined as p=2 using a number of information criteria.

Table 12. Causality Test Results

H_0	χ^2	Prob.	Decision
LOGICT3 does Granger-cause LOGSE	6.136701	0.0465	logICT3→logSE
All does Granger-cause LOGSE	24.34792	0.0005	All → logSE
LOGSE does't Granger-cause LOGICT3	0.461584	0.7939	logSE ··· logICT3

All does Granger-cause LOGICT3	5.776997	0.4486	All ··· logICT3
A → B means causality runs from A to B.			
A ··· B means no causality between A and B			

Table 12 indicates that in the model where the dependent variable is total service exports, there is causality from mobile phone subscriptions, an ICT indicator, to service exports, and there is also causality from all variables to service exports. In the model where the dependent variable is ICT3, it is observed that there is no causality from service exports to ICT3, and there is no causality from all variables to ICT. In other words, there is a unidirectional causality between service exports and ICT3 in Armenia; while ICT influences service exports, service exports do not influence ICT3.

4.2.1. Diagnostic Tests

Breusch-Godfrey Test was used to test the existence of autocorrelation problem in the series. It is seen that the probability values in Table 13 are greater than 0.05.

H_0: *There is no autocorrelation.*

H_1: *There is autocorrelation.*

In this case, based on the hypothesis " H_0 = *There is no autocorrelation*", no autocorrelation problem is found in the model up to three lags in all three country groups. At 5% significance level, the hypothesis is accepted and it is decided that there is no autocorrelation problem in the analysis.

Table 13. Diagnostic Test Results

Diagnostic test results for Azerbaijan				
Lag	**VAR Autocorrelation LM test results**		**VAR Variable variance problem**	
1	17.10356	0.3789	Ki-Kare	Olasılık değeri
2	15.81703	0.4658		
3	20.55995	0.1961	182.7510	0.1051
Diagnostic test results for Georgia				
1	15.61667	0.4800	Ki-Kare	Olasılık değeri
2	8.231093	0.9417		
3	19.06692	0.2652	176.5079	0.1762
Diagnostic test results for Armenia				
1	30.74254	0.0145	Ki-Kare	Olasılık değeri
2	22.27137	0.1346		
3	18.30567	0.3063	156.0231	0.3514

For the detection of the problem of varying variance, which is the last stage of satisfying the stability condition

Breusch-Pagan-Godfrey Test was used.

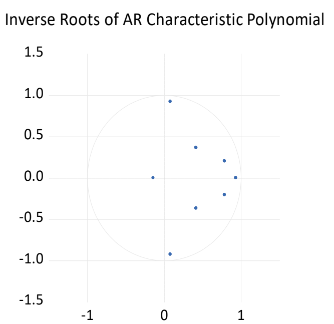
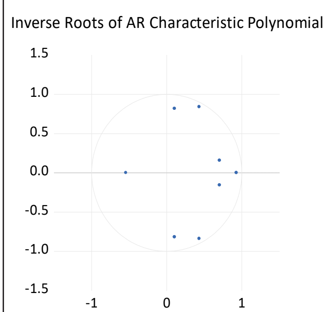
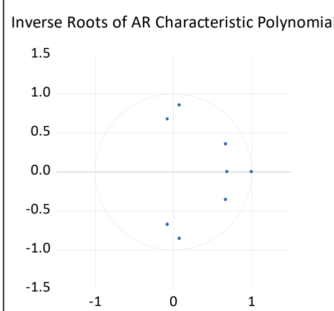
H_0: *There is no variance.*

H_1: *There is variable variance.*

According to the varying variance test results given in Table 13, the hypothesis is accepted at 5% significance level and therefore, it is determined that there is no varying variance problem in the analysis.

Table 14 shows the characteristic inverse roots of the model. It is seen that all inverse roots for all countries remain within the unit circle and satisfy the stability condition.

Table 14. Stability Condition Table

<table>
<tr><th>AZERBAJIAN</th><th>GEORGIA</th><th>ARMENIA</th></tr>
<tr>
<td>
</td>
<td>
</td>
<td>
</td>
</tr>
</table>

5. Discussion, Conclusion and Policy Recommendations

In this study, which examines the impact of information and communication technologies on services exports for the period 2003-2022, the ICT indicator is used separately for each country, taking into account differences in internet infrastructure, internet access and internet usability. The literature summarized in Table 2 (Wang & Choi, 2019; Lichy et al., 2022; Nasir & Kalirajan, 2016; Freund & Weinhold, 2002) suggests that country income, technology utilization level, technology skills, and technology ownership have a positive and significant effect on trade in services. On the other hand, in the same literature table, it is observed that service exports (Freund & Weinhold, 2002; Choi, 2010; Nath & Liu, 2017) are negatively affected by population. The exchange rate (Lichy et al., 2022) and the level of employment in the service sector (Wang & Choi, 2019) are also very important in trade in services. This study analyzes the relationship between service exports and information and communication technologies in the context of the service industry in the South Caucasus using time series data.

In the study, it was analysed that fixed telephone subscription, which is the variable included as an ICT indicator, causes service exports in Azerbaijan, service exports cause individual internet usage in Georgia, and mobile telephone subscription causes service exports in Armenia. In all three countries, a unidirectional causality relationship was found between ICT and services exports. Factors affecting exports have always been one of the most researched topics both in terms of literature and governments. Although there are no direct empirical studies on the subject, based on studies examining the relationship between ICT and trade, this study, which investigates the role of ICT in service exports in the South Caucasus, concludes that the relationship between mobile phone subscriptions as an ICT indicator and service exports in Armenia is in line with Nasir and Kalirajan (2016), Wardani et al. (2020), Choi (2010), and Freund & Weinhold (2002). On the other hand, individual internet use, which yields the best results as an ICT indicator for Georgia, is found to be effective in service exports as in (Wang and Choi, 2019; Lichy et al., 2022). It can be said that the result obtained for Azerbaijan is similar to (Nath & Liu, 2017; Wang & Choi, 2019).

In the late 20th century, ICT, which started to develop and spread rapidly, marked the first step in stimulating international trade and the global economy. It is crucial for service exports, which play a significant role in increasing foreign exchange earnings, creating new employment opportunities, and generally contributing to economic development. The findings of this research make important contributions to the development of service trade in the South Caucasus. It is emphasized that prioritizing ICT is essential to develop service exports in the South Caucasus countries, enabling them to reach the level of developed countries in the service sector and even surpassing these developed countries. This is because, as in many other sectors, prioritizing ICT in the service sector enhances industrial productivity and global competitiveness performance. Moreover, ICT is frequently preferred by many institutions, organizations, and individuals for its advantages such as time and cost savings. Therefore, the importance of knowledge-based digital production in the economy is increasing. In the South Caucasus countries, especially in Azerbaijan, fixed telephone subscriptions were identified as the reason for service exports, while in Armenia, mobile telephone subscriptions were identified as the key factors in facilitating and developing service trade.

In the study, the ICT variables that have the most significant impact on service sector exports are individual internet usage and mobile phone subscriptions. At this juncture, it is crucial to ensure that other ICT indicators are also effective in countries. This would facilitate the increase in service exports through the influence of ICT. Encouragement should be provided to ICT-related departments in existing universities, and high-quality software developers should be trained at reputable universities. ICT initiatives and advancements in service sector areas such as accounting, finance, and law should be prioritized, enabling the transformation of ideas into successful businesses. This is because the development of ICT enhances communication possibilities, reduces transaction costs, and thus can positively affect international trade by promoting greater commercial participation in the foreign trade process.

As internet coverage in Caucasian countries increases, it is imperative to expand, improve, and strengthen internet infrastructures. Otherwise, the digital divide is inevitable. In the study, the ICT variables that have the most significant impact on service sector exports are individual internet usage and mobile phone subscriptions. At this point, it is of great importance to ensure that other ICT indicators are also effective in countries. This will facilitate the increase in service exports with the impact of ICT. As internet coverage increases in Caucasian countries, it is imperative to expand, improve and strengthen internet infrastructures. In Azerbaijan in particular, efforts should be made to reduce income inequality, and improvement and development policies should be established to reduce economic inequalities and poverty, thus preventing the digital divide caused by insufficient internet access for a significant part of the population. Public-private partnerships should be established to extend broadband internet access to rural areas in Azerbaijan. Data centres in accordance with international standards should be established in major cities such as Baku and investments should be made for the security and sustainability of these centres. ICT related departments in existing universities should be encouraged and high quality software developers should be trained in reputable universities. The curricula of ICT and software engineering programmes in universities should be updated and emphasis should be placed on practical skills. ICT initiatives and advancements in service sector areas such as accounting, finance and law should be prioritised to transform ideas into successful businesses. This is because the development of ICT increases communication possibilities, reduces transaction costs and thus can positively affect international trade by encouraging greater business participation in the foreign trade process. Technology incubation centres should be established in Baku and other major cities, providing financing and mentoring services for start-ups. Special incentive programmes should be created for R&D projects in the ICT field. New trade agreements should be negotiated to incentivise exports of digital services. Tax reductions and incentives should be provided for companies operating in the ICT sector. Effective political and macroeconomic management should prioritise the use of technology in all service areas, especially internet infrastructure. Robotic coding education should be integrated into pre-school education and ICT education should be encouraged. Otherwise, ICT facilities alone will not be sufficient and service sector exports will not reach the desired level.

Since there is a risk of a digital divide in Georgia and Armenia, as in many other countries, necessary incentives and support should be provided for R&D activities in the Caucasus countries and universal access to ICT should be ensured for all individuals at minimal or no cost. This would facilitate the development of various service sectors and thus increase service exports. Infrastructure investments should be made for the rapid adoption of 5G technology in Armenia. The establishment of 5G networks through public-private partnerships should be encouraged and the regulations on this issue should be updated rapidly. Venture capital funds should be established with the participation of local and international investors. Innovation competitions and hackathons should be organised regularly to encourage innovative ideas. Patent and intellectual property rights laws should be strengthened to protect ICT innovations. Grant and support programmes should be established for ICT projects. Microfinance and microcredit programmes should be initiated to meet the financing needs of small-scale enterprises. Regional cooperation with neighbouring countries in the field of ICT should be developed. High-speed internet access should be expanded in Georgia, especially in Tbilisi and other big cities, and should be extended to rural areas. Local cloud service providers should be supported and international co-operation should be encouraged. Support the development of the entrepreneurship ecosystem by establishing technology parks and incubation centres in Tbilisi. E-commerce and data protection laws should be updated and regulations should be made to ensure the security of digital services. Special state support programmes should be established for the ICT sector and innovative projects should be financed.

Investments should be made to expand high-speed internet access in rural and urban areas to help the South Caucasus countries strengthen their ICT sector and increase competitiveness in services exports. The development of digital infrastructure such as data centres, cloud computing and secure internet networks should be encouraged. ICT-focused education programmes in schools and universities should be established and updated, tax reductions and subsidies should be provided for research and development activities in ICT, innovation centres and incubation programmes should be established for start-ups and technology companies. Legal arrangements should be made to facilitate and protect the export of digital services, and tax incentives should be applied for ICT investments and companies exporting services. Free trade agreements that facilitate the export of ICT services should be concluded. Marketing campaigns should be organised to ensure that the country is internationally recognised for ICT services. Co-operation between the public and private sectors for ICT services and innovation projects should be encouraged. In addition, knowledge sharing and co-operation in ICT fields between developing countries and Caucasian countries can be encouraged.

In future studies, the impact of ICT in specific service sectors such as health, education, finance and tourism can be analysed separately. The effects of ICT use in these sectors on exports can be analysed. Furthermore, research can be conducted on analysing the potential target markets for service exports of the South Caucasus countries and developing entry strategies for these markets. Collaboration models in the field of ICT among the South Caucasus countries and the impact of these collaborations on service exports can be analysed. The impact of ICT infrastructure (internet penetration, broadband access, mobile connections) on service exports can be analysed by making a cross-country comparative analysis of how the development of ICT infrastructure affects service exports in different countries. The contribution of the use of leading digital platforms (freelancing sites, e-commerce platforms) to service exports can be analysed on a sectoral basis to see how digital service platforms increase service exports and in which sectors they are most effective. How small and medium-sized enterprises (SMEs) can increase their service exports by using ICT can be analysed through surveys and case studies among SMEs. Laws and regulations governing the export of digital services, how they can be harmonised with ICT and how this harmonisation affects exports can be determined by comparing digital service export regulations in different countries. The impact of ICT on service exports can be analysed by comparing regions with different levels of development.

The most important constraint in carrying out this study was the lack of reliable and up-to-date data.

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RESEARCH ARTICLE

ASSESSING THE DISTRIBUTIVE EFFECTS OF MINIMUM WAGE: EVIDENCE FROM TURKEY

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Abstract

This study explores the effects of the 2016 minimum wage hike in Turkey on wage distribution up to 2022 by using a difference-in-differences methodology. This approach employs unconditional quantile regressions by utilizing variation in the bite of the minimum wage across NUTS2 regions in Turkey and utilizes data from the Turkish Household Labor Force Survey (HLSF). The findings indicate that the 2016 minimum wage increase positively affects wages in the lower quantiles while having a negative impact on wages in the higher quantiles. Consequently, this leads to a wage compression effect, ultimately resulting in a reduction in wage inequality, as supported by descriptive analysis.

Keywords: Minimum wage, Inequality, Wages

JEL codes: J31, J38

1. Introduction

The impact of minimum wage policies on different worker groups has been a widely debated topic in the economic literature, with thorough reviews conducted by numerous researchers, including Card and Krueger (1995), Brown (1999), Machin and Manning (1997), Rubery (2003), Manning (2011), Neumark and Wascher (2008), and Levin-Waldman (2018). The minimum wage can potentially reduce wage inequality due to its diverse impact across the wage distribution. When the minimum wage increases, the lowest earners witness a significant boost in their income, while middle-income individuals experience a moderate gain and high earners encounter only a minimal or negligible rise. Simultaneously, employment and overall economic output experience only a slight decline as workers tend to shift to more productive firms (Engbom and Moser, 2022). Research from developed countries (; Butcher et al. 2012; Machin and Manning 1994; Teulings 2003; Card and Krueger 1994; DiNardo et al. 1996; Lee 1999; Dickens and Manning 2004; Stewart 2012; Autor et al. 2008; Autor et al. 2016, Caliendo et al. 2017; Fortin and Lemieux 2000; Vandekerckhove et al. 2018; Bossler and Shank 2023) showed that minimum wages have a substantial impact on reducing wage inequality. In developing countries, the effect of minimum

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wages on wage inequality is often unclear due to segmented labor markets and weak enforcement. However, an increasing body of research, such as the studies conducted by Bosch and Manacorda (2010) in Mexico, Lin and Yun (2016) in China, Engbom and Moser (2022) and Sotomayor (2021) in Brazil, Khurana et al. (2023) in India, and Lombardo et al. (2022) in Latin American countries, has demonstrated that raising minimum wages has the potential to result in a decrease in wage inequality in emerging economies.

Turkey exhibits numerous similarities with other developing economies, such as having a large informal sector and increasing reliance on informal employment, weak labor institutions lacking effective law enforcement, an economy with low allocation efficiency characterized by a significant presence of small, low-productivity firms, and low productivity of labor attributed to insufficient levels of human capital among employees. Consequently, a substantial portion of the workforce comprises low-wage employees who receive minimum wage compensation, a characteristic commonly observed in many developing-country economies (Bossavie et al., 2019). Hence, examining Turkey's case broadens our currently limited understanding of how minimum wages impact wage inequality in developing countries characterized by these labor market features. The distributional effects of Turkey's 2016 minimum wage increase demand special attention because it was driven mainly by exogenous political competition ahead of the 2015 elections rather than being a consequence of the economy's internal dynamics. Following this increase, In the period following 2016, the Turkish economy witnessed a notable decline in institutional autonomy driven by increased authoritarianism. This era has also seen the revival of reactive ad-hoc policies, macroeconomic instability, and the challenging external conditions stemming from the post-2015 US Federal Reserve tightening period. In conjunction with rising inflation, this economic landscape prompted the government to adopt a proactive minimum wage policy, leading to a rise in the real minimum wage, surpassing the average real wages after 2016.

In recent years, considerable attention has been dedicated to studying the effects of the minimum wage on Turkey's labor market, with particular attention given to employment outcomes rather than its impact on wage inequality. However, the available evidence in the literature concerning labor market outcomes of minimum wage in Turkey is notably varied. Güven et al. (2011) found no relationship between the minimum wage and employment in the Turkish manufacturing industry throughout 1969-2008. Pelek (2015) investigated whether the national minimum wage has influenced employment rates of workers aged 15-29 by taking regional disparities into account and found no disemployment effect for this age group. Gürcihan-Yüncüler and Yüncüler (2016) explored the consequences of the 2004 minimum wage increase on labor market outcomes, discovering a favorable impact on working hours and informality. Dağlıoğlu and Bakır (2015) revealed a positive correlation between the minimum wage and employment, showing distinct effects on men and women. Aslan (2019) demonstrated a reduction in informality within Turkey's market attributable to the minimum wage increases between 2003 and 2017. Notwithstanding these findings, several studies have identified adverse impacts of the minimum wage on employment in Turkey. Öztürk (2007) investigated the effect of the minimum wage on the Turkish labor market before the 2000s and found that it has a detrimental impact on employment among low-productivity workers and the number of part-time jobs. Papps (2012) examined the effect of the 2004 minimum wage increase on employment, revealing a decline in the likelihood of formal employees retaining their jobs. Bakis et al. (2015) demonstrated that the minimum wage hike in 2004 motivated teenagers to pursue schooling and decreased their participation in the labor force. Bossevie et al. (2019) showed that the 2016 minimum wage increase resulted in a significant rise in the destruction of formal firms, especially small ones characterized by low productivity levels, ultimately reducing the total number of formal enterprises in the economy. Gürsel et al. (2018) discovered a strong positive relationship between the minimum wage rise in 2016 and the proportion of informal employment within the labor market.

While the number of studies examining the influence of minimum wage on wage inequality in Turkey is limited, their findings are notably consistent, suggesting that the minimum wage has a decreasing effect on wage inequality. Pelek (2013) examined the impact of the 2004 minimum wage increase on wages by decomposing the wage differences and trends in wage dispersion prior to and afterward the rise and found that the minimum wage has been instrumental in diminishing wage inequality among male and female wage earners. Bakis and Polat (2015) examined the evolution of wage inequality using the decomposition approach between 2002 and 2010 and showed that the 2004 minimum wage increase played a significant role in decreasing wage inequality. Bakis and Polat (2015) conducted a decomposition analysis to examine the changes in wage inequality from 2002 to 2010. Their findings highlighted that the 2004 minimum wage increase played a significant role in decreasing wage inequality. Ekşi and Kırdar (2015) attributed the decline in wage inequality from 2002 to 2011 to the 2004 minimum wage increase. Gürcihan-Yüncüler and Yüncüler (2016) demonstrated that the 2004 minimum wage increase contributed to a reduction in wage inequality using a quasi-experimental approach. Tamkoç and Torul (2020) explored the role of minimum wage hikes in reducing wage inequality by conducting a counterfactual analysis. Işık et al. (2020) showed the positive effects of the 2016 minimum wage increase on wages of most demographic groups and informal workers by employing a difference-in-differences approach. Bakış and Polat (2023) revealed that the minimum wage increases in 2004 and 2016 contributed to reducing the wage disparity between the upper and lower percentiles by employing a decomposition approach.

The results of this study are consistent with the limited prior research on the effects of the 2016 minimum wage increase on wage disparity, indicating a reduction in wage inequality. This research examines the impact of the 2016 minimum wage increase along the wage distribution up to 2022 using a difference in differences (DID) approach, applying unconditional quantile regressions to data from the Turkish Household Labor Force Survey (HLSF). The study reveals that the 2016 minimum wage increase positively impacts wages in the lower quantiles while it negatively impacts wages in the higher quantiles. This results in a wage compression effect, ultimately leading to an improvement in wage dispersion, as corroborated by descriptive analysis. As far as our knowledge extends, the primary contribution of this study is its pioneering attempt to investigate the impact of the 2016 minimum wage increase on different quantiles of the wage distribution while extending the analysis until 2022.

The rest of the study is structured as follows: Section 2 offers insights into the Institutional framework and the Minimum Wage Increase in 2016. Section 3 examines the evolution of minimum wage and wage Inequality in Turkey between 2004 and 2022. Section 4 describes the data and difference-in-differences methodology employed in the study. Section 5 presents the empirical results, and Section 6 concludes.

2. The Institutional Framework and the Minimum Wage Increase in 2016

Figure 1 provides an overview of the historical evolution of Turkey's minimum wage legislation. Despite the introduction of modern minimum wage legislation in Turkey in 1936, the onset of adverse conditions resulting from the Second World War delayed the implementation of minimum wage until the 1950s. The minimum wage was determined by regional commissions between 1951 and 1967. Considering local or regional characteristics in determining the minimum wage has not been adequately reflected in practice. Since there has not been sufficient harmony between local minimum wage commissions, significant minimum wage differences have occurred between very close or distant regions within the same period, even in regions with similar economic and social structures. A central committee was instituted to determine minimum wage rates for specific sectors and regions in 1967. From 1969 to 1973, minimum wage rates were exclusively set for the industrial sector. Subsequently, in 1973, a separate minimum wage was introduced for individuals working in agriculture and forestry. Between 1969 and 1974, the central minimum wage was determined regionally, and the national minimum wage was officially introduced in 1974. In 1989, the practice of varying minimum wage rates between industry, agriculture, and forestry was abandoned. Since then, the country has consistently maintained a uniform minimum wage nationwide without distinctions based on regions or sectors (Yolvermez, 2020, 244).

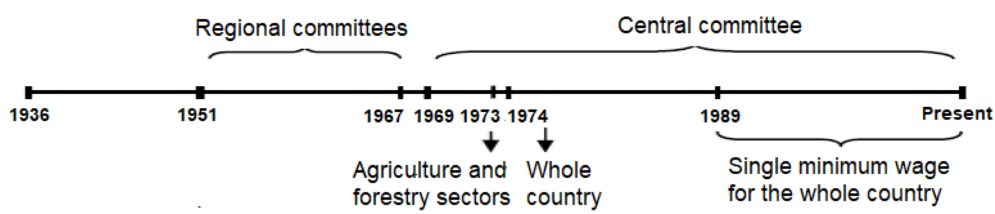

Figure 1. Historical Evolution of Minimum Wage Legislation in Turkey
Source: Yılmaz-Eser and Terzi (2008, 131)

As specified in the Minimum Wage Law of 2004, the central committee consists of 15 members, including two representatives from the Ministry of Labor and Social Security, one representative from the Turkish Statistical Institute (Turkstat), one representative from the Undersecretariat of Treasury, one representative from the Ministry of Development, five representatives to be elected by the highest labor organization that has the largest number of workers for different branches of business, and five representatives to be elected by the employer organization that has the largest number of employers for other branches of industry. Although the current law stipulates a minimum wage determination period of at least every two years, the minimum wage has been adjusted annually or biannually over the past two decades. Between 2004 and 2015, the committee disclosed separate minimum wage rates for the initial and latter halves of the upcoming year each December. This decision, taking inflation expectations into account, meant that the revised minimum wage rates would be enforced on January 1 and July 1 of the following year. Between 2016 and 2021, the committee established a single minimum wage level for the entire year. However, in 2022 and 2023, due to elevated inflation rates, the committee reverted to setting the minimum wage biannually.

A minimum wage is regarded as setting an external wage floor within labor markets. It is often considered as a form of collective bargaining, particularly in countries with significantly limited labor union representation and weak labor market institutions, such as Turkey. This approach directly or indirectly influences a significant portion of the workforce (Kahveci and Pelek, 2021). Isık et al. (2020) highlighted three crucial characteristics of the Turkish labor market that should be considered when assessing the impact of the minimum wage on labor market outcomes. These characteristics, including high informality levels, gender-based disparities in labor force participation, unemployment, wages favoring men, and regional disparities in unemployment and labor force participation rates, collectively intensify the influence of the minimum wage in labor markets. Furthermore, the low level of unionization and reduced coverage of collective bargaining agreements in Turkey make the minimum wage a primary indicator in the labor market. In 2019, Turkey's collective bargaining coverage was only 8.5%, significantly lower than the OECD average of 32.1% for the same year.

During the 2000s, the most significant hike in minimum wage occurred in 2016, resulting in a 33% nominal and 25% real increase. In contrast, previous increases in the minimum wage were characterized by gradual and smaller changes, typically falling within the range of approximately 5% to 8% in nominal terms. This significant increase in the minimum wage was primarily a result of the electoral competition in 2015. Before the November 2015 elections, all political parties committed to significantly raising the minimum wage as part of their campaign promises, engaging in a competitive stance to offer the most substantial increase. Consequently, following the election, the national minimum wage was established at 1300 TL on January 1, 2016. The emergence of the new minimum wage through the political process indicates that the change was primarily influenced by external political factors rather than internal economic dynamics. The significant magnitude and the externally driven nature of the 2016 minimum wage increase provide a robust experimental setting for investigating the causal effects of the increase on wage distribution.

3. Evolution of Minimum Wage and Wage Inequality in Turkey

Figure 2 depicts the trends in the share of employees earning the minimum wage and below-average real wages and real minimum wages. While real minimum wages and average real wages exhibited similar trends up until 2016, a substantial divergence between the two has become evident in subsequent years. After the minimum wage increase in 2016, the growth in the real minimum wage has consistently surpassed the growth in the average real wage. This can be attributed to two main factors: the growing share of workers paid minimum wage and below and the depreciation of average nominal wages relative to inflation. As illustrated in Figure 1, the percentage of employees earning minimum wage and below has consistently remained high, averaging 28.6% between 2004 and 2015 and notably increasing to an average of 41.1% between 2016 and 2022. The highest share was observed in 2019, reaching 46.25%. In contrast, the percentage of workers receiving wages at or below the minimum wage in the European Union (EU-27) for the same year was 15% (ILO, 2021). These indicators from Figure 2 alone paint a picture of Turkey becoming a nation of minimum wage earners.

Figure 2. Evolution of Real Wage, Real Minimum Wage, Indicators for Minimum Wage

Source: Author's calculations based on data from HLSF and Ministry of Labor and Social Security of Turkey.
Note: Real wage and Below Employment Ratio are calculated using HLSF data

In Figure 3, the variance of log wages and the log differences of wage percentiles are used to analyze the overall wage disparity between 2004 and 2022, which is assessed through the log wage variance. It seems that wage inequality initially decreased until the global crisis of 2008, but then it rose until 2010. Subsequently, it diminished until the end of the period, with the decreasing trend briefly interrupted by a slight increase between 2016 and 2020. Panel b of Figure 2 presents the wage inequality measured by the log differences of various percentiles. The wage disparity at the higher end of the distribution, exemplified by the gap between the 90th and the 50th percentile, experienced an increase following the 2008 global financial crisis. This increase stabilized until 2016 when a decline was observed; from then on, it remained relatively constant.

Dispersion in the lower half of the wage distribution, as depicted by the gaps between p50 and p20, p50 and p15, and p50 and p5, demonstrates a declining trend over the given period. The measures in Figure 3 show that wage inequality at the end of the period is lower than at the beginning, except for log differences of p90-p50, which have almost the same value at the two endpoints of the analysis period.

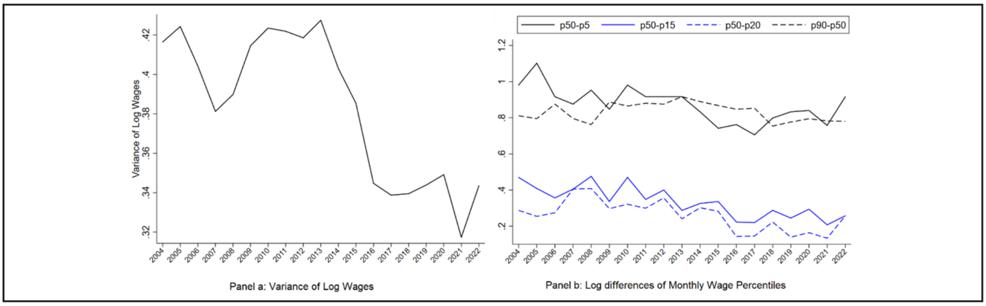

Figure 3. Evolution of the Wage Disparity

Source: Author's calculations based on HLSF data.

As shown in Figure 4, the proportion of the minimum wage to the median wage in Turkey is significant and exhibits an upward trend. Since 2007, it has demonstrated an upward trajectory, reaching a peak of 89% by 2022. This ratio stands at 54.7% for full-time workers across 32 OECD countries. The percentage of the minimum wage to mean wage exhibits a similar trend, although it is smaller than the percentage of the minimum wage to the median wage. The period between 2009 and 2015 shows a relatively stable trend for the minimum wage to mean wage ratio, followed by a sharp increase in 2016 and a gradual rise until the period's end. In 2022, it reached its highest point at 0.71%, while the average for 32 OECD countries was 43.2% for the same year. Despite being lower than the ratio of minimum wage to average wage, the percentage of minimum to 90th percentile wage follows a similar trend.

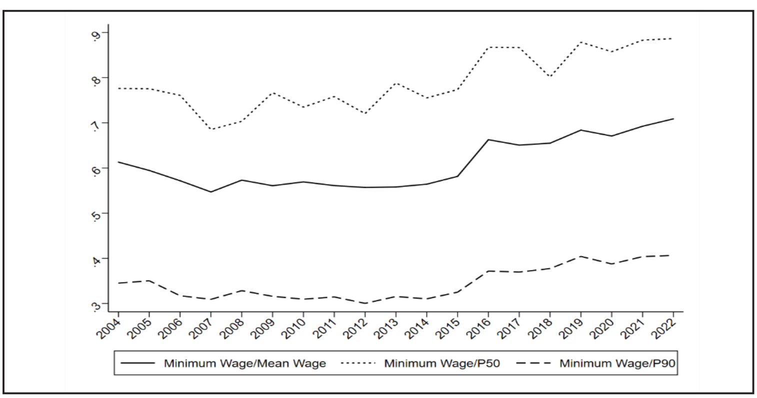

Figure 4. Minimum Wage Statistics for Turkey

Source: Author's calculations based on data from HLSF and the Ministry of Labor and Social Security of Turkey.

Kernel density plots are utilized to understand better the impact of the 2016 minimum wage increase on wage distribution, as they are practical tools for visualizing wage levels and highlighting the parts affected by the minimum wage changes. Figure 5 displays the Kernel estimates of the real monthly wages of wage earners in 2015 and 2016. As expected, wage distribution exhibits a right skew, with the mean being higher than the median each year. The minimum wage truncates the wage earners' distribution, resulting in spikes at the minimum wage level for both years. The most notable change from 2015 to 2016 is a leftward shift of the lower end of the distribution, while the right segment has sustained a relatively stable pattern. In other words, a significant wage increase occurred at the bottom of the wage distribution. In contrast, the wages at the top of the wage distribution remained substantially stable between the two years.

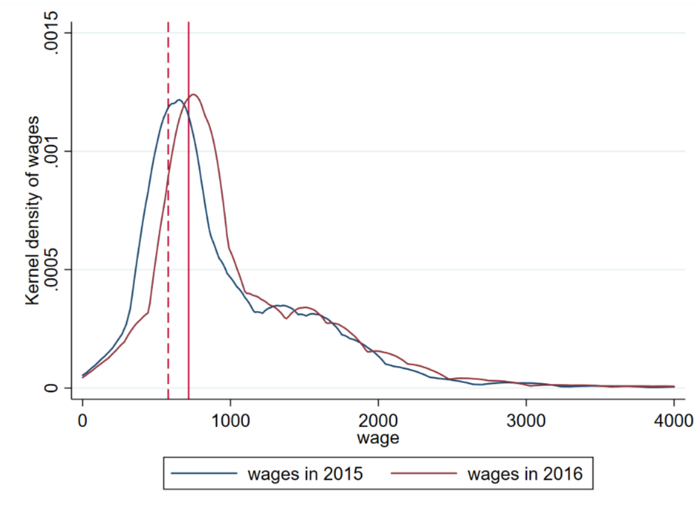

Figure 5. Real Monthly Wage Distributions in 2015 and 2016

Source: Author's calculations based on data from HLSF

Notes: Vertical dash line and solid line indicate the real minimum wages in 2015 (578.5 TL) and 2016 (716.4 TL), respectively.

The HLFS, 2003 and 2005, own calculations

Table 1 presents an alternative perspective on the influence of the minimum wage, showing growth in the real monthly wage (calculated as the log-point differences) over various wage distribution bins. This approach follows the methodology outlined by Stewart (2012) and Bossler and Schank (2023). The 2015-2016 period stands out significantly from the previous periods due to the magnitude of the growth rate in each bin. From 2015 to 2016, the growth of each wage bin was at least ten times larger than the growth rates observed in previous periods, highlighting the significant impact of the 2016 minimum wage increase on the wage distribution. Yet, wage growth increases slightly before the actual change, which could be attributed to an upward wage movement at the lower deciles or a minor anticipation effect.

The first row of Table 1 shows that individuals initially located at the very low end of the wage distribution undergo the most significant wage growth. More robust wage growth at the lowest segment of the distribution is expected due to the concept of mean reversion in wages, which suggests that workers with lower incomes can experience more substantial wage growth than those with higher incomes. From 2015 to 2016, individuals receiving wages between 100% and 110% of the minimum wage also saw notable wage growth. However, it was 1.6 times smaller in magnitude than the wage growth observed in the 'below minimum wage' group. While this effect diminishes when moving up the wage distribution (row 2 and beyond), significant increases in wage growth can be observed across the wage distribution. This suggests an upward wage shift attributed to the 2016 minimum wage increase, extending to individuals not directly impacted by the change.

Table 1. Wage Growth by Wage Distribution Bins

	(1)	(2)	(3)	(4)
	wage growth in the periods			
Initial Wage Bin	2012-2013	2013-2014	2014-2015	2015-2016
below minimum wage	-0.056	0.044	0.051	0.479
100% - 110% minimum wage	0.015	0.016	0.031	0.284
110% - 120% minimum wage	0.030	0.020	0.030	0.257
120% - 130% minimum wage	0.029	0.007	0.028	0.237
130% - 140% minimum wage	0.016	0.010	0.016	0.215
140% - 150% minimum wage	0.020	0.026	0.017	0.208
≥ 200% minimum wage	0.043	-0.017	-0.006	0.181

Source: Author's calculations based on HLSF data.

4. Data and Methodology

The data employed in this study is individual-level cross-sectional data from the Household Labor Force Survey (HLFS) conducted by the Turkish Statistical Institute (TurkStat) covering 2004 to 2022. The HLFS complies with the definitions and concept standards established by The European Union Statistics Office (Eurostat) and collects data on various aspects of the labor force structure in Turkey, including wage, economic activity, occupation, employment status, working hours, as well as information on the duration of unemployment and the type of occupation sought by the unemployed. The HLSF covers the non-institutional population, a minimum of 366,000 households per year between 2004 and 2022. The sample in this study covers employees aged between 15 and 65. It intentionally incorporates all employees with different statuses, including part-time and temporary, without imposing restrictions based on gender to ensure that the groups that the minimum wage is most likely to have a significant influence on are not excluded from the analysis.

The wage variable in the data is the monthly wage earned from the individual's main job activity, including all wage supplements. The monthly wage is adjusted for inflation with the base year set as 2008. In the study, the OLS (Ordinary Least Squares) and quantile regression analyses utilize the natural logarithm of real monthly wages. Additionally, regional variations are taken into account at the NUTS2 level, consisting of 26 statistical regions as defined and published by Turkstat.

The impact of the January 2016 minimum wage increase on different parts of the wage distribution is investigated using the difference in differences (DID) specification employing unconditional quantile regressions, which utilize regional differences in the bites of the rise in the minimum wage. Defining y_{it} as the natural logarithm (log) of the monthly wages for an individual i at the time t, Firpo et al. (2009, 2018) formulated The RIF (re-centered influence function) of y_{it} for various deciles τ, and the variance of y_{it}, σ^2 as follows:

$$RIF\ (y_{it}, \tau) = y_\tau - \frac{\tau - I[y \leq = y_\tau]}{f_Y(y_\tau)} \qquad (1)$$

$$RIF(y_{it}, \sigma^2) = (y_{it} - \mu)^2 \qquad (2)$$

In a linear regression context, Firpo et al. (2009) explained that employing the RIF of y_{it} as the dependent variable leads to the generation of unconditional quantile regression. The coefficients derived from the RIF regressions are the average marginal impact on y_{it} at the specified percentile, τ.. The definition of "bite" holds significant importance in this analysis, so the literature explores various alternatives. The most notable is the Kaitz index, the ratio between the minimum and regional average wages. A greater Kaitz value suggests that the minimum wage has a more significant effect. However, it's worth noting that changes in the Kaitz index are not solely driven by shifts resulting from the minimum wage; fluctuations in other segments of the wage distribution also influence this indicator. Another bite measure, "fraction," focuses on the percentage of workers directly impacted by minimum wage increases. It illustrates how much the minimum wage impacts the eligible working population by showing how many of them are affected by the change. A higher proportion of employees earning less than the minimum wage prior to its rise indicates a significant number of workers whom the minimum wage change will impact. Several studies such as Card (1992), Stewart (2002), Dolton et al. (2015), Caliendo et al. (2018), Bossler and Schank (2023), and Wittbrodt (2022) employ the "fraction" bite measure, which is calculated as the proportion of the employed individuals receiving wages less than the minimum wage. Bossler and Schank (2023) explain the primary benefit of utilizing regional variation as its ability to capture spillover effects caused by adjustments in the minimum wage in a specific region. For instance, if one employee experiences a wage increase while some other's wage is reduced in remuneration, the overall wage effect within the labor market remains neutral, regardless of which of the two employees is being considered. This study calculates the bites as the proportion of individuals paid below the minimum wage level before the increase in 2016 in 26 NUTS2 statistical subregions of Turkey. Figure 6 depicts the "bite measure" variation across NUTS2 regions in Turkey, highlighting a diverse impact among different geographical areas. However, the minimum wage increase has the most pronounced effect on the southern-east region of Turkey.

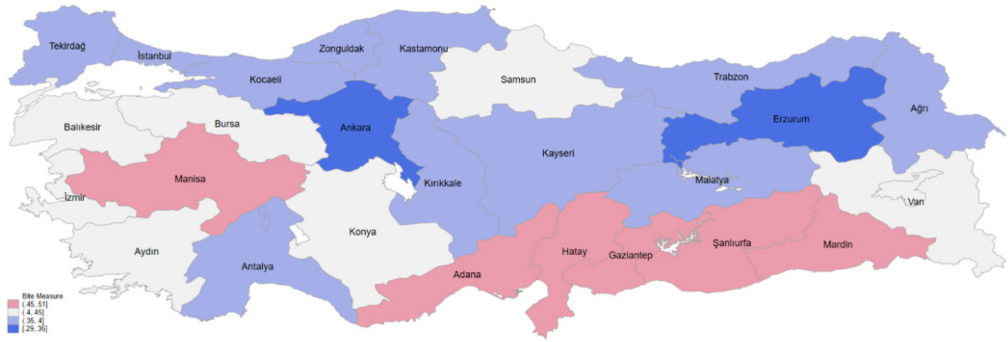

Figure 6. Dispersion of The Bites Across NUTS2 Regions in Turkey

Source: Author's calculations based on HLSF data.

The following difference-in-differences specification from Bossler and Schank (2023) was used in this study:

$$RIF(y_{it},\tau) = \phi * Bite_r + \pi * Bite_r * Trend_t + \sum_{t=2015}^{2022} \delta_t * Bite_r * Year_t + \sum_{t=2014}^{2022} \lambda_t * Year_t + +\epsilon_{it}, \quad (3)$$

where y_{it} represents the natural logarithm (log) of the monthly wages for an individual i at the time t. The dependent variable of the specification is the RIF of y_{it} computed for various deciles of the distribution, τ, the variance of y_{it}, σ^2. Equation 3 comprises treatment effect interaction terms for the years after the 2016 increase in the minimum wage, as well as for 2015, to capture any anticipatory effects. The specification also includes common time effects, the effect of the bite capturing constant level differences over the years, and an interaction between a time trend and the bite variable to capture an existing bite-specific trend. This specification makes it possible to identify the average impact of the minimum wage bite at various points along the unconditional distribution of log monthly wages following the change in the minimum wage policy in 2016. Standard errors are calculated with bootstrap (50 replications) and clustered at the regional level.

An important consideration regarding the empirical specification of this study is to decide whether to include a bite-specific trend ($Bite_r * Trend_t$) in the model. Although t-tests justify the presence of the interaction term in the specification, a graphical examination is conducted to examine how the coefficients' trends differ between models that include the bite-specific trend and those that do not, in line with the approach outlined by Bossler and Schank (2023). Following unadjusted DID specification which does not control for the term ($Bite_r * Trend_t$) is estimated:

$$RIF(y_{it},\tau,unadjusted) = \phi * Bite_r + \sum_{t=2014}^{2022} \delta_t * Bite_r * Year_t + \sum_{t=2014}^{2022} \lambda_t * Year_t + +\epsilon_{it}, \quad (4)$$

Figure 7 displays the estimated δ_t values along the wage distribution for the unadjusted and trend-adjusted models. The unadjusted treatment effects graph (panel A) shows notably positive time trends, particularly evident at the lower end of the wage distribution. This suggests a more pronounced rise in low wages within regions experiencing higher bite measures compared to low wages in areas with lower minimum wage impacts. The trend-adjusted graph (panel b) reveals a suppressed trend at the bottom and middle of the wage distribution, especially after the minimum wage increase in 2016. On the other hand, the high end of the wage distribution (70th and 90th percentiles) has a more pronounced negative trend, particularly evident after 2016. This graphical examination justifies the bite-specific trend along with the statistically significant t-test results for the bite-specific trend in the specification.

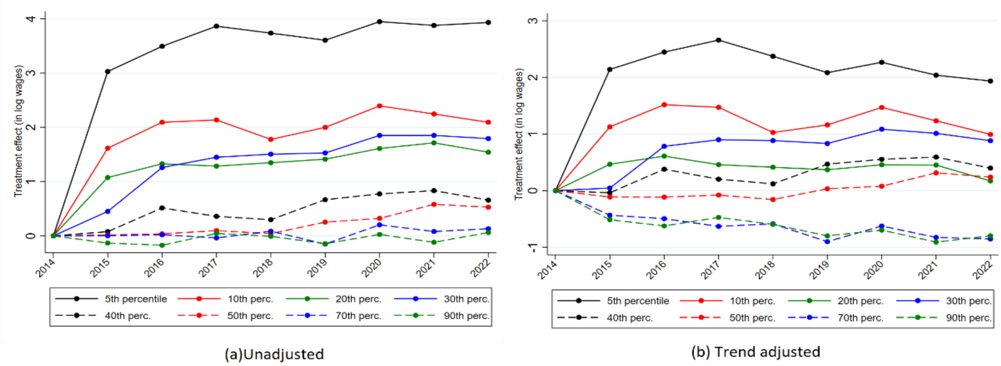

Figure 7. Treatment Effects (coefficients of interaction terms for bite and each year)
Source: Author's calculations based on HLSF data.
Note: The reference year is 2014

5. Findings

The primary results for Equation 3 are displayed in Table 2. Column (1) shows the OLS results, and the presence of statistically insignificant interaction terms suggests the necessity to investigate the effects of the minimum wage increase across the wage distribution. The following columns provide details on the effects of the minimum wage increase on the (RIF) of log wages, covering the 5th to 90th percentiles. At the 5th percentile, the size of the treatment effect interactions ranges from 2 to 2.7, indicating that a 10% increase in the bite results in a minimum 20% increase in monthly wages. The interaction term for 2015 exhibits a less pronounced positive effect than 2016, which can be interpreted as an indicator of anticipation. The wage effect increases to 2.45 in 2016, when the minimum wage significantly increased, suggesting that a 10% increase in the regional bite causes a 24.45% increase in wages at the 5th percentile. In 2017, the impact of the bite reaches its highest level, with a 26.6% increase in response to a 10% increase in the regional bite. In the subsequent years, the magnitude of the effect gradually diminishes, reaching 20.4% in 2021, and it attains statistical significance in 2022. At the 10th percentile, the magnitude of wage effects is approximately half of what is observed at the 5th percentile, ranging between 1 and 1.48. Like the situation observed at the 5th percentile, a significant interaction term for 2015 indicates anticipation of the minimum wage increase in 2016. A 10 percent increase in the regional bite results in an 11% rise in the wage level in the 10th percentile in 2015, while the wage effect increases to 15.2% in 2016. Following the minimum wage increase, the wage effect of the regional bite experienced a decline in the two subsequent years, namely 2017 and 2018, then followed a similar trend to that observed at the 5th percentile. No statistically significant wage effect is attributable to the regional bite at the 20th percentile. However, positive wage effects can be observed at the 30th and 40th percentiles with no significant anticipation effects. Examining treatment effect interactions for each year, the magnitudes of the significant wage effects diminish from the 5th to the 40th percentiles. The impact of minimum wage bites is not statistically significant at the 50th and 60th percentiles, but adverse effects are evident in the higher segments of the wage distribution. In 2017, following the minimum wage increase, the negative wage impact of the bite becomes statistically significant at the 70th percentile. Throughout 2022, the adverse impact varies between 6 % and 8.5 % for every 10 % increase in the minimum wage bite. At the 90th percentile, the anticipation effect is evident, resulting in a 5 % wage decrease in 2016. the adverse wage impact of the regional bite stands at 6%, but it decreases to 4.7% in the subsequent year. The negative wage effect of a 10% increase in the regional bite ranges between 4.7 % and 9% at the 90th percentile.

The RIF quantile results, which show that the regional bite has a positive wage impact at the lower percentiles and a negative wage impact at the higher percentiles, imply that the minimum wage increase in 2016 decreased wage dispersion. Column (11) presents the effect of the 2016 minimum wage increase on the variance of log wages. The anticipation effect is observable in 2015 as a 6% decrease in the variance of log wages. In the year of the minimum wage increase and the subsequent years, the treatment effect is about 10%. This suggests that a 10 percent increase in the regional bite results in a corresponding 10% reduction in the variance of log wages. The decrease in the variance of log wages due to an increase in the regional bite aligns with the RIF quantile findings, which indicate a reduction in wage inequality resulting from the minimum wage increase in 2016. These findings are supported by the declining trend in the variance of log wages, as depicted in Figure 3. To assess the robustness of the observed reduction in inequality resulting from the minimum wage increase in 2016, a sensitivity analysis regarding the sample definition is conducted, as outlined in Bossler and Schank (2023).

Table 2. The "DID" Results for The Impact of the 2016 Minimum Wage Increase

Dependent Variable	(1) Ln(w)	(2) RIF($\tau_{5\%}$)	(3) RIF($\tau_{10\%}$)	(4) RIF($\tau_{20\%}$)	(5) RIF($\tau_{30\%}$)	(6) RIF($\tau_{40\%}$)	(7) RIF($\tau_{50\%}$)	(8) RIF($\tau_{60\%}$)	(9) RIF($\tau_{70\%}$)	(10) RIF($\tau_{90\%}$)	(11) RIF(σ^2)
Explanatory Variables											
Bite	-1.733***	-4.424***	-2.757***	-1.866***	-2.392***	-1.343***	-1.241***	-1.436***	-1.240***	-0.699***	1.016***
	(0.203)	(1.322)	(0.794)	(0.348)	(0.341)	(0.069)	(0.112)	(0.254)	(0.313)	(0.244)	(0.332)
Bite*trend	0.076***	0.158***	0.087***	0.109***	0.072***	0.021***	0.023***	0.045***	0.078***	0.068***	-0.022**
	(0.003)	(0.026)	(0.024)	(0.021)	(0.009)	(0.005)	(0.009)	(0.009)	(0.009)	(0.006)	(0.010)
D2014	-0.029***	-0.078**	-0.007	0.021	0.006	0.002	-0.007	0.012	-0.052***	-0.035	0.009
	(0.009)	(0.037)	(0.027)	(0.044)	(0.034)	(0.014)	(0.016)	(0.013)	(0.016)	(0.024)	(0.012)
D2015	-0.054	-0.907**	-0.460**	-0.162	-0.071	0.077	0.053	-0.011	0.107	0.166***	0.260***
	(0.070)	(0.354)	(0.191)	(0.107)	(0.138)	(0.064)	(0.098)	(0.121)	(0.107)	(0.060)	(0.098)
D2016	-0.012	-1.101***	-0.595***	-0.224**	0.006	0.177**	0.137	0.165	0.087	0.203***	0.415***
	(0.059)	(0.419)	(0.211)	(0.105)	(0.117)	(0.069)	(0.136)	(0.109)	(0.183)	(0.072)	(0.109)
D2017	-0.067	-1.201***	-0.614***	-0.224**	-0.027	0.039	0.091	0.157	0.135	0.089	0.385***
	(0.049)	(0.400)	(0.216)	(0.112)	(0.123)	(0.079)	(0.106)	(0.121)	(0.126)	(0.079)	(0.122)
D2018	-0.094	-1.181**	-0.545**	-0.250**	-0.072	0.037	0.105	0.060	0.070	0.089	0.393***
	(0.061)	(0.491)	(0.221)	(0.120)	(0.138)	(0.075)	(0.096)	(0.075)	(0.137)	(0.062)	(0.131)
D2019	-0.064	-1.125**	-0.583**	-0.258**	-0.086	0.124**	0.056	0.206*	0.274	0.213**	0.399***
	(0.077)	(0.442)	(0.232)	(0.113)	(0.127)	(0.058)	(0.112)	(0.109)	(0.176)	(0.095)	(0.103)
D2020	-0.096	-1.226***	-0.714***	-0.339***	-0.174	0.083	0.131	0.145**	0.154**	0.183*	0.469***
	0.089)	(0.443)	(0.244)	(0.112)	(0.123)	(0.051)	(0.100)	(0.067)	(0.077)	(0.104)	(0.101)
D2021	-0.101	-1.202***	-0.657***	-0.354***	-0.154	0.073	0.138	0.172**	0.148	0.186**	0.514***
	(0.069)	0.439)	(0.238)	(0.113)	(0.114)	(0.052)	80.127)	(0.076)	(0.152)	(0.073)	(0.105)
D2022	-0.134**	-1.273***	-0.636***	-0.330***	-0.177	-0.005	0.068	0.100	0.142	0.122**	0.492***
	(0.061)	(0.438)	(0.224)	(0.115)	(0.112)	(0.048)	(0.072)	(0.085)	(0.116)	(0.054)	(0.114)
Bite*D2015	0.036	2.141**	1.127**	0.466	0.046	-0.038	-0.112	-0.150	-0.434	-0.514***	-0.655**
	(0.188)	(0.885)	(0.467)	(0.327)	(0.367)	(0.174)	(0.275)	(0.258)	(0.267)	(0.128)	(0.256)
Bite*D2016	0.122	2.447**	1.517***	0.610	0.783**	0.379*	-0.114	-0.370	-0.495	-0.623***	-1.092***
	(0.153)	(1.082)	(0.534)	(0.371)	(0.332)	(0.230)	(0.360)	(0.287)	(0.364)	(0.184)	(0.284)
Bite*D2017	0.159	2.659***	1.472***	0.459	0.899**	0.202	-0.078	-0.407	-0.631**	-0.473***	-1.015***
	(0.111)	(0.999)	(0.527)	(0.395)	(0.354)	(0.210)	(0.288)	(0.294)	(0.289)	(0.180)	(0.315)
Bite*D2018	0.091	2.373*	1.027*	0.413	0.884**	0.120	-0.157	-0.318	-0.587**	-0.595***	-1.024***
	(0.141)	(1.232)	(0.591)	(0.409)	(0.377)	(0.214)	(0.279)	(0.211)	(0.261)	(0.157)	(0.349)
Bite*D2019	0.062	2.084*	1.161**	0.369	0.833**	0.468***	0.033	-0.434	-0.901***	-0.799***	-0.979***
	(0.181)	(1.126)	(0.542)	(0.432)	(0.368)	(0.165)	(0.334)	(0.283)	(0.346)	(0.210)	(0.306)
Bite*D2020	0.178	2.267*	1.470**	0.457	1.085***	0.554***	0.078	-0.236	-0.625***	-0.697***	-1.086***
	(0.203)	(1.178)	(0.604)	(0.427)	(0.345)	(0.140)	(0.321)	(0.176)	(0.136)	(0.238)	(0.299)
Bite*D2021	0.116	2.039*	1.233**	0.452	1.012***	0.593***	0.314	-0.357*	-0.826***	-0.909***	-1.257***

	(0.154)	(1.173)	(0.599)	(0.462)	(0.343)	(0.145)	(0.302)	(0.216)	(0.236)	(0.193)	(0.316)
Bite*D2022	0.020	1.936	0.993*	0.171	0.882**	0.398***	0.239	-0.322	-0.853***	-0.798***	-1.146***
	(0.154)	(1.176)	(0.507)	(0.459)	(0.340)	(0.139)	(0.231)	(0.250)	(0.220)	(0.151)	(0.348)
Constant	7.380***	7.514***	7.176***	7.023***	7.223***	6.983***	7.088***	7.350***	7.537***	7.774***	0.000
	(0.088)	(0.542)	(0.313)	(0.129)	(0.137)	(0.030)	(0.052)	(0.106)	(0.144)	(0.098)	(0.135)
Observations	1,770,236	1,770,236	1,770,236	1,770,236	1,770,236	1,770,236	1,770,236	1,770,236	1,770,236	1,770,236	1,770,236
Cluster	26	26	26	26	26	26	26	26	26	26	26

Source: Author's calculations based on HLSF data.

Notes: Bootstrap clustered robust standard errors are reported in parentheses (clustered at NUTS2 regional levels). *p<0.1,**p<0.05,***p<0.01. Data is limited to employees between 15 and 65 years old.

Table 3. The "DID" Results for The Impact of the 2016 Minimum Wage Increase Within – Restricted Samples

	(1)	(2)	(3)	(4)	(5)
Dependent Variable	RIF(σ^2)	RIF(σ^2)	RIF(σ^2)	RIF(σ^2)	RIF(σ^2)
Explanatory Variables					
Bite	1.016***	0.743**	0.642**	0.430*	-0.012
	(0.332)	(0.311)	(0.313)	(0.238)	(0.153)
Bite*trend	-0.022**	-0.018***	-0.024***	-0.010	0.006
	(0.010)	(0.006)	(0.006)	(0.006)	(0.004)
D2014	0.009	0.009	0.010	0.000	0.002
	(0.012)	(0.009)	(0.006)	(0.008)	(0.007)
D2015	0.260***	0.261***	0.256***	0.206***	0.151**
	(0.098)	(0.093)	(0.090)	(0.074)	(0.063)
D2016	0.415***	0.364***	0.349***	0.303***	0.265**
	(0.109)	(0.081)	(0.087)	(0.061)	(0.049)
D2017	0.385***	0.340***	0.304***	0.255***	0.207***
	(0.122)	(0.110)	(0.111)	(0.091)	(0.075)
D2018	0.393***	0.328***	0.281**	0.244***	0.198**
	(0.131)	(0.115)	(0.112)	(0.087)	(0.083)
D2019	0.399***	0.291***	0.277***	0.228***	0.188***
	(0.103)	(0.089)	(0.080)	(0.065)	(0.070)
D2020	0.469***	0.400	0.401***	0.357***	0.274***
	(0.101)	(0.084)	(0.097)	(0.070)	(0.054)
D2021	0.514***	0.478***	0.441***	0.390***	0.303***
	(0.105)	(0.092)	(0.088)	(0.087)	(0.070)
D2022	0.492***	0.463***	0.402***	0.351***	0.266***
	(0.114)	(0.111)	(0.100)	(0.088)	(0.079)
Bite*D2015	-0.655**	-0.647***	-0.627***	-0.543***	-0.400***
	(0.256)	(0.233)	(0.224)	(0.190)	(0.151)
Bite*D2016	-1.092***	-0.934***	-0.876***	-0.795***	-0.690***
	(0.284)	(0.203)	(0.221)	(0.159)	(0.122)
Bite*D2017	-1.015***	-0.877***	-0.764***	-0.688***	-0.569***
	(0.315)	(0.270)	(0.273)	(0.224)	(0.183)
Bite*D2018	-1.024***	-0.840***	-0.695**	-0.670***	-0.580***
	(0.349)	(0.286)	(0.282)	(0.220)	(0.208)
Bite*D2019	-0.979***	-0.674***	-0.611***	-0.576***	-0.522***
	(0.306)	(0.234)	(0.210)	(0.173)	(0.179)
Bite*D2020	-1.086***	-0.839***	-0.814***	-0.808***	-0.656***
	(0.299)	(0.217)	(0.239)	(0.164)	(0.119)
Bite*D2021	-1.257***	-1.097***	-0.951***	-0.960***	-0.788***
	(0.316)	(0.241)	(0.226)	(0.227)	(0.175)
Bite*D2022	-1.146***	-1.024***	-0.827***	-0.853***	-0.700
	(0.348)	(0.284)	(0.255)	(0.226)	(0.201)
Constant	0.000	0.060***	0.068	0.153	0.306***
	(0.135)	0.126)	(0.130)	(0.097)	(0.065)

Restrictions					
Males Only		Yes	Yes	Yes	Yes
Full-Time Only			Yes	Yes	Yes
Permanent Jobs Only				Yes	Yes
Prime Age Only					Yes
Observations	1,770,236	1,261,707	1,293,112	1,147,717	950,589

Source: Author's calculations based on HLSF data

Notes: Bootstrap clustered robust standard errors are reported in parentheses (clustered at NUTS2 regional levels) *p<0.1,**,p<0.05,***,p<0.01.

Considering that much of the prior literature has primarily concentrated on full-time male workers within their prime working age range, typically between 25 and 55 years, a stepwise approach in limiting the sample to encompass these specific demographics is followed. This approach is designed to systematically assess the applicability of the above findings within the well-established framework of existing literature. Given that prime-age male employees in full-time and permanent positions are less likely to be significantly impacted by changes in the minimum wage, narrowing the sample to include these groups exclusively may not yield results consistent with the baseline findings of this study.

Hence, this robustness check could uncover how much the 2016 minimum wage change influenced the Turkish labor market. Table 3 displays the corresponding results derived from the restricted samples, the dependent variable of which is the variance of log wages. Column (1) presents the primary results from Table 2, while the subsequent columns show the results obtained by applying sample constraints incrementally. Findings from the restricted samples consistently indicate that the adverse effect of the 2016 minimum wage increase on wage inequality persists, albeit gradually decreasing in magnitude after introducing additional sample restrictions.

6. Conclusion

This study examines the impact of the significant minimum wage increase in 2016 on wage distribution in Turkey from difference-in-differences estimation, taking into account the variations in the minimum wage bite across Turkey's NUTS2 regions. This specification applied to the unconditional wage distribution of real monthly wages using data from the HLFS. The dataset encompasses all employees, regardless of their employment status, including part-time and temporary workers, both men and women. This approach ensures that the analysis includes groups most likely to be significantly affected by the minimum wage. The bite of the minimum wage is calculated as the fraction of workers paid below the minimum wage level before the increase in 2016 in 26 NUTS2 statistical subregions of Turkey. The graphical examination of the bite measure across the regions reveals a diverse impact among NUTS2 regions, highlighting that the minimum wage rise in 2016 has the most pronounced effect on the southern-east part of Turkey. Prior to the difference-in-differences estimation, a series of descriptive analyses is undertaken to understand the impact of the minimum wage rise in 2016 on the wage distribution. The comparison of the Kernel estimates of the real monthly wages of wage earners in 2015 and 2016 reveals a significant wage increase occurred at the bottom of the wage distribution while the top of the wage distribution remained substantially stable between the two years. As an alternative way to examine the change in the wage distribution, growth rates in the real monthly wage across different distribution bins are calculated. Examining wage bin growth reveals that the 2015-2016 period differs significantly from previous periods due to the magnitude of the growth rate in each bin, with diminishing growth rates as one moves up the wage distribution, which is consistent with the Kernel estimates comparisons of 2016 and 2016. Additional descriptive analyses of the evolution of the minimum wage and wage inequality from 2004 to 2022 reveal that, following the minimum wage increase in 2016, the growth in the real minimum wage has consistently surpassed the growth in the average real wage, which can be attributed to two main factors: the growing share of workers paid minimum wage and below and the depreciation of average nominal wages relative to inflation. In other words, since 2016, there has been a convergence of average real wages toward the real minimum wage in Turkey, coinciding with the increasing number of minimum wage earners. This trend reflects the growing influence of the minimum wage in labor markets, further accentuating deteriorating macroeconomic conditions since 2016. The results of the difference-in-differences analysis indicate that the 2016 minimum wage increase positively affected wages in the lower quantiles while negatively impacting wages in the higher quantiles. This led to a wage compression effect, resulting in reduced wage inequality. This conclusion is supported by a significant decrease in the variance of log wages, which dropped by approximately 10 to 12 percent annually after the introduction of the minimum wage until 2022, in addition to the descriptive analyses conducted in the study. Robustness checks, including restrictions by gender, age, and employment status, confirmed the enduring impact of the minimum wage increase on reducing wage inequality. This study's results align with the findings of the limited previous literature on the distributive effects of the 2016 minimum wage rise in Turkey. This study contributes to the existing literature as a pioneering attempt to investigate the impact of the 2016 minimum wage increase on various quantiles of wage distribution while extending the analysis up to 2022.

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RESEARCH ARTICLE

STOCHASTIC CONVERGENCE OF INCOME IN TURKIYE: A METHODOLOGICAL REINVESTIGATION OF PROVINCES

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Abstract

This study revisits income convergence among Turkish provinces for 1992-2019 and differs from most empirical literature due to its unique structural and methodological framework. Stochastic convergence is tested by employing a battery of panel stationarity tests that allow cross-sectional dependence and structural breaks. Breaks are further analyzed with respect to the nature of breaks as sharp and smooth. Sharp breaks are identified endogenously, while smooth breaks are accounted for using the Fournier approximation. Although σ-convergence is detected, there are no shreds of evidence of stochastic convergence at the panel level. Univariate test statistics demonstrate that at the provincial level, there is no single case that applies to all provinces. As additional dimensions of the data-generating process are evaluated in the testing procedure, outcomes about stochastic convergence slightly shift for provinces. However, findings at the panel level remain consistent and do not produce stochastic convergence. At the provincial level, mixed results are obtained.

Keywords: Stochastic Convergence, Fourier Approximation, Panel Unit Root, Regional Economics, Stationarity

JEL Classification: C23, O47, R10

1. Introduction

The existence of regional disparities and their patterns are quite crucial not just from an academic intellectual curiosity viewpoint but also because they have the power to govern the agenda of policy-makers. In that respect, this study tries to revisit some old yet still relevant issues in Turkiye using a province level. The first and foremost aim is to explore convergence structure employing a solid methodological approach quite different from the common practice in the literature.

The idea of convergence, in the contemporaneous understanding, was introduced by Solow (1956) under the framework of the neoclassical growth theory, which is inevitable under the diminishing return to physical or human capital assumptions because that tenet forces each economy[1] to approach its own steady state in the long run. Relative distance to their steady states

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1 Hereinafter, instead of economies or countries, "regions" are used in order to be aligned with the content of this study.

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governs their growth rate, producing *conditional* convergence. However, once the model posits identical preferences and homogenous technologies for all regions, they share the same unique steady state irrespective of initial conditions. Regions far away from the long-run equilibrium grow faster than regions closer, and eventually; poorer regions become as rich as the initially rich regions. That sort of catch-up is called *absolute* convergence. The neoclassical growth theory does not predict absolute convergence but occurs as a particular case.

On the other side, endogenous growth theories initiated by Romer (1986) and Lucas (1988) criticized the critical building blocks of the neoclassical growth theory and incorporated positive externalities or spillovers into the setup through increasing returns into the production function in the form of intentional human capital accumulation and R&D activities. The long-run growth is determined within the model endogenously rather than by taking it as an exogenous factor. This strand of growth literature predicts no convergence or even *divergence* as the initial condition of a region is determined by endogenous drivers. Besides absolute and conditional convergence, Galor (1996) proposes a third alternative: club convergence – regions with similar structural features (e.g., initial conditions) or heterogeneity in factor endowments form clusters with distinct steady-states even in the neoclassical growth model.

Empirical testing of convergence can be classified broadly into four different methodologies [2]: i) cross-section approach, ii) panel approach, iii) times series approach, and iv) distribution approach (Islam, 2003). The distribution approach fundamentally differs from the rest because it deals with the entire income distribution instead of directly working with regression analysis. Markov chain analysis is one way to account for such distribution dynamics (Quah, 1993a). The other tool is σ-convergence, a convergence that seeks a decline in income dispersion and is quantified by either standard deviation or coefficient of variation (Dowrick and Nguyen, 1989; Friedman, 1992; Boyle and McCarty, 1997). The cross-section approach (Barro, 1991; Mankiw, Romer, and Weil, 1992; Barro and Sala-i Martin, 1992) searches for a negative relationship between initial income level and growth rate of per capita income. This method is called β-convergence. Absolute or conditional β-convergence division is based on whether other structural characteristics beyond the initial income level are controlled. However, such initial level regression (i.e., Barro-type regressions) is criticized by Quah (1993b) for being an example of Galton's fallacy. Additionally, differences in initial technology levels are seen in the error term of the regression, and besides the capital deepening as a source of income convergence, technology diffusion as the other source disappears due to the assumption of homogenous technologies across regions (Islam, 2003). The panel approach (Islam, 1995; Caselli, Esquivel, and Lefort, 1996; Barro, 1996) is viewed as a potential candidate for solving this problem. Explicit control of technology terms has a dual advantage. First, the technology term captures more than technology (e.g., other aspects of the economic structure), and second, omitted variable bias stemming from unobserved heterogeneity is solved with the individual [3] (regional) effect in the regression equation (Temple, 1999; Islam,

2 For an extensive literature review on different conceptualizations of convergence phenomenon, please see Temple (1999), Islam (2003) and Durlauf, Johnson and Temple (2005).

3 Caselli, Esquivel and Lefort (1996) stress that for regions that share similar technologies, bias from individual effect

2003). The time series approach to convergence underpins this study. Thus, I discuss it in-depth in the Section 3.

The plan of the paper is as follows. Section 2 presents the related empirical literature on Turkiye. Section 3 provides the theoretical foundations of the stochastic convergence. Section 4 introduces the data and some descriptive analysis. Section 5 outlines the econometrics methodology and Section 6 concludes.

2. Related Literature Review

Several studies were conducted to reveal convergence dynamics in Turkiye's regions. Filiztekin (2018) proved the existence of conditional β-convergence among 65 provinces between 1975 and 1995, yet divergence was detected via σ-convergence, particularly after the late 1970s to late 1980s. Tansel and Güngör (1999) studied the same period for 67 provinces using productivity instead of regional GDP. They found that absolute β-convergence and the speed of convergence accelerated after 1980, which is attributed to the liberalization practices. However, σ-convergence exhibited different patterns for western and eastern provinces. Although, Doğruel and Doğruel (2003) documented β-convergence for 1987-1999 for all, high-income and low-income provinces, failure to find σ-convergence in low-income provinces was tied to decreased public investments in those regions. Karaca (2004) investigated the 1975-2000 period, but could not find evidence of β-convergence, and divergence was explored as income dispersion widened. Önder, Deliktaş, and Karadağ (2010) conducted a series of panel techniques and observed conditional convergence for NUTS-2 regions during 1980-2001; however, the transportation component of public capital stock was found as a factor that exacerbated regional disparities. Gömleksiz, Şahbaz, and Mercan (2017) also supported the role of government in stimulating convergence for 2004-2014 at NUTS2 level.

Using the panel approach, Bolkol (2019) found shreds of evidence on both unconditional and conditional convergence for different regional units, including provinces, from 2005 to 2017. Despite the fall in the variation, strong arguments about σ-convergence were unavailable, but the 2008-09 crisis period contributed to the convergence experience. Later, Bolkol (2023) added an endogenous growth perspective and stressed that policies based on R&D personnel would not lead to convergence but rather a divergence.

Aksoy, Taştan, and Kama (2019) observed convergence clubs rather than absolute or conditional convergence for the 1987-2001 and 2004-2017 periods. Similar convergence clubs were obtained in Sakarya, Baran, and İpek (2024) for 2004-2022. However, two subsets, 2004-2016 and 2017-2022, exhibited different patterns. The tendency of convergence turned into divergence for 81 provinces. There are also some studies (Gezici and Hewings, 2004; Aldan and Gaygısız, 2006) mainly concentrating on spatial links and some studies (Aldan and Gaygısız, 2006; Karahasan,

may be trivial. Time-specific effect captures world growth and commons shocks (Durlauf et al. 2005).

2017 and 2020) with Markov chain analysis; yet all of them demonstrated the continuity in the regional income variation. Beside the β-convergence, a strand of literature was flourished after Carlino and Mills (1993), Quah (1993a), Bernard and Durlauf (1995).

Erlat and Özkan (2006) used CADF panel unit root and tested the time series approach to convergence in Turkish provinces. They found that different regions involved different patterns signaling some sort of club formations but failed to get clear evidence on absolute convergence for 1975-2000. Aslan and Kula (2011) analyzed 67 provinces from 1975 to 2001 with a univariate LM unit root test that enabled the endogenous determination of structural breaks. Allowing two structural breaks resulted in stochastic convergence for all provinces except Bitlis, Erzurum, and Hakkari so that shocks to relative income had only transitory impact. Durusu-Çiftçi and Nazlıoğlu (2019) applied a series of univariate unit root tests to 73 provinces from 1992 to 2013, allowing for sharp shifts and smooth shifts. However, they took the presence of stochastic convergence as a necessary but not sufficient condition and checked β-convergence for each province following Tomljanovich and Vogelsang (2002). The clear divergence between eastern and western provinces was reached. Akkay (2022) employed similar univariate unit root tests as Durusu-Çiftçi and Nazlıoğlu (2019) and extended the terminal year to 2019. All provinces experienced stochastic convergence, and this result remained consistent regardless of whether structural breaks, primarily in 2002 and 2008, were taken into account.

The literature on regional stochastic convergence in various countries is extensive. Notable studies include Tomljanovich and Vogelsang (2002) on regions in the United States, DeJuan and Tomljanovich (2005) on Canadian provinces, Constantini and Arbia (2006) on Italian regions, Carrion-i-Silvestre and German-Soto (2009) on Mexican regions, and Misra, Kar, Nazlıoğlu, and Karul (2024) on Indian states.

3. Theoretical Foundations of Stochastic Convergence

Quah (1992) encapsulates the convergence phenomenon using several approaches and defines one approach as the absence of unit root or deterministic time trend in income disparities between countries that is intrinsically and fundamentally different from initial level regression analysis. Bernard and Durlauf (1995; 1996) also express that regions [4] i and j convergence between time t and $t+T$, when the per capita output difference is expected to fall. $y_{i,t}$ corresponds to natural logarithm of real per capita output and if $y_{i,t} > y_{j,t}$ then the previous statement can be demonstrated as $E\left(y_{i,t+T} - y_{j,t+T} | I_t\right) < y_{i,t} - y_{j,t}$ in the time series context. This structure is later elaborated to capture variant aspects of the convergence such that two regions are said to converge if the long-term forecasts of per capita output for both regions are equal to a fixed time t, conditional on some information set at t, including time, current and deeper lags of $y_{i,t}$ (see

4 Definitions are based on countries but since this study explores regional convergence, from now on "region" replaces "country" in such definitions.

Bernard and Durlauf, 1995 and 1996). The benchmark unit appears as a problem to be solved, and there are two paths of practice: choosing a reference country or taking a sample average [5].

According to Evans and Karras (1996) i regions are said to convergence if, and only if, a common trend a_t, which is unobservable by nature and equivalent to technology [6], and finite parameters $\mu_1, \mu_2, \ldots, \mu_i$ exist such that

$$\lim_{T \Rightarrow \infty} E\left(y_{i,t+T} - a_{t+T} | I_t\right) = \mu_i \tag{1}$$

μ_i is a parameter governing the balanced growth path of the region i. Common trend is obtained by averaging over i regions so that

$$\lim_{T \Rightarrow \infty} E(\bar{y}_{t+T} - a_{t+T} | I_t) = \frac{1}{i} \sum_{i=1}^{i} \mu_i \tag{2}$$

where $\bar{y}_t = \sum_{i=1}^{i} \frac{y_{i,t}}{i}$. The level of common trend is defined as $\lim_{T \Rightarrow \infty} E(\bar{y}_{t+T} - a_{t+T} | I_t) = 0$, so common trend equals to average behavior of i economies. To eliminate it, we subtract (2) from (1) and generate

$$\lim_{T \Rightarrow \infty} E\left(y_{i,t+T} - \bar{y}_{t+T} | I_t\right) = \mu_i \tag{3}$$ [7]

(3) is isolated from common trend and is left with deviations of per capita income from average behavior. Namely, long run forecasts of relative per capita incomes approach to a constant as the forecasting horizon tends to infinity and this can be directly tested by checking the stationarity of the deviation of output, $y_{i,t+T} - \bar{y}_{t+T}$ (Evans and Karras, 1996; Bernard and Durlauf, 1995 and 1996) [8].

Using a similar rationale, Carlino and Mills (1993) first suggest *(stochastic) convergence* to test whether shocks to relative income are temporary or not. In case of stationarity, idiosyncratic regional specific factors are also immune to long-run economic growth and shocks have only transitory impacts (Carrion-i Silvestre and Soto, 2009). On the other hand, non-stationarity triggers a shock of permanent deviations in relative per capita income and hampers any tendency of stochastic convergence. Thus, future trajectories of such behaviors cannot be projected. Temple (1999) also emphasizes the link between convergence and stationarity testing but is aware of how hard to get precise interpretations.

A body of empirical literature on this issue emerges in the context of whether or not there is a time trend [9]. Trend stationarity case is named as *stochastic convergence* (Carlino and Mills, 1993;

5 Latter strategy is adopted to bring into alignment with regional convergence literature. See Islam (2003) for possible problems of taking deviations from either reference economy or sample average.

6 "Not just technology but resource endowments, climate, institutions and so on; it may therefore differ across countries" Mankiw, Romer and Weil (1992: 5-6).

7 Bernard and Durlauf (1995; 1996) used $\lim_{T \Rightarrow \infty} E\left(y_{i,t+T} - \bar{y}_{t+T} | I_t\right) = 0$ version of the formula.

8 For the bivariate case, incomes have to be cointegrated. See Bernard and Durlauf (1995); Stengos and Yazgan (2014) for details.

9 See Islam (2003) for discussion of stochastic and deterministic trends.

Strazicich, Lee and Day, 2004) or *catching-up* (Oxley and Greasley, 1995; Cunada and Garcia, 2006) while level stationarity as either *deterministic convergence* [10] (Li and Papell, 1999; Cunada and Garcia, 2006) *or long-run convergence* (Oxley and Greasley, 1995). However, Li and Papell (1999) remark a caveat about a time trend as it can cause permanent per capita income differences making it vulnerable to criticism. Zero mean stationarity, without a constant and time trend case, is also discussed (Bernard and Durlauf, 1995; Cunada and Garcia, 2006). A generic explanation of divergence, in our case, is that per capita income gap between a region and country average consistently widens and requires non-stationarity.

However, it is worth noting that the time series approach, to a large extent, is inherently statistical and not linked explicitly to growth theories because initial conditions have no role in the long-run trajectories (Oxley and Greasley, 1995; and Durlauf, Johnson, and Temple, 2005). On the other hand, the impacts of initial cross-country differences in physical and human capital on the long-run patterns construct the backbone of neoclassical and endogenous growth theories (see Durlauf, *et al.*, 2005). Evans and Karras (1996) and Evans (1998) put some effort into reconciling the time series approach with growth theories, aiming at strengthening the weak ties. Evans (1998) argues that $y_{i,t+T}$ reverts to common trend, measured by $\bar{y}_{t+T}$, lends some support to exogenous growth theory, while the case of non-reverting $y_{i,t+T}$ to common trend provides what the endogenous growth models require [11]. The former case corresponds to stationarity, whereas non-stationarity leads to the latter case. Relevant models need to be tested appropriately to get more definitive and concrete outcomes (Oxley and Greasley, 1995), so this study avoids such certain claims. A further taxonomy is also possible rooted in Evans and Karras (1996) by modifying equation (3) as follows: i) *absolute convergence* takes place when $\mu_i = 0$ for all is, or ii) *conditional convergence* if $\mu_i \neq 0$ for some i. To be clearer based on the distinction made above, zero mean stationarity implies the same steady-state for all regions (King and Ramlogan-Dobson, 2014) and is analogous to *absolute convergence* (Cunada and Garcia, 2006). It is also proposed that a constant term (Strazicich *et al.* 2004) or a_{t+T} can capture some time-invariant differences giving rise to *conditional convergence* (Islam, 2003). As a matter of fact, most of the earlier literature is based upon Dickey-Fuller type equation estimation (Carlino and Mills, 1993; Oxley and Greasley, 1995; Bernard and Durlauf, 1995; Li and Papell, 1999). Using a well-behaved neoclassical production function, the following equation [12] can be written to test for convergence

$$y_{it} = \mu_i - \beta g t + (1+\beta) y_{i,t-1} + \varepsilon_{it} \qquad (4)$$

If region subscripts are removed, it becomes the Dickey-Fuller equation [13] with constant and time trend. To achieve (stochastic) convergence, $(1+\beta)$ has to be less than one, that is to say β should be negative or it should not contain unit root (Islam, 2003). Although technology (A_t) specification

10 Li and Papell (1999) label Bernard and Durlauf (1995; 1996) case as *deterministic convergence.*

11 For a more straightforward interpretation, cross-section specific intercepts should be checked as well. For more, see Evans and Karras (1996), and Evans (1998).

12 The proof of this equation can be found in Islam (1995 and 2003).

13 Dickey and Fuller (1979) model (c) is $y_t = \mu - \beta t + \rho y_{t-1} + \varepsilon_t$

plays a role in the type [14] of convergence, this study quests for only stochastic convergence under different sets of assumptions of the data-generating process (DGP).

Bernard and Durlauf (1995 and 1996) put forward a prominent remark about the inappropriateness of such time-series [15] testing for economies that are far from their steady-states, pointing out that unit root null hypothesis can be spuriously accepted because, in this case, the data may be generated by a transitional law of motion rather than by an invariant random process. Thus, the sample moments of the data are not representative to the population moments. This research acknowledges the aforementioned empirical concerns. Even though the true DGP for provincial per capita income in Turkiye may be difficult to have or even unattainable fully because provinces may not be close to their steady-states, the true DGP can be approximated considering all probable and relevant peculiarities of the data.

4. Data and Descriptive Analysis

The income per capita relative to a benchmark, mostly the average of the regions, is needed to test the stochastic convergence. Nevertheless, the fact that per capita income is not reported regularly at the provincial level prevents the use of official statistics retrieved from Turkstat. The official series covers 1987-2001 (with the *base* year 1987) and 2004-2022 (with the *reference* [16] year 2009). Methodological change to the chain-volume index [17] from the constant-price approach in the calculation of GDP and the missing period of 2002-2003 do not make it feasible (Düşündere, 2019; Akkay, 2022). Thus, in unreliable [18] or unavailable subnational data, luminosity can be used as a proxy for economic performance (Chen and Nordhaus, 2011; Henderson, Storeygard, and Weil, 2012). For this purpose, Düşündere (2019 and 2020) estimates luminosity-based income per capita at the provincial level for 1992-2019 [19] using satellite nighttime light data. This study utilizes that new dataset and converts provincial GDP (chain-volume index) in Turkish Lira into GDP per capita for 81 provinces using population data. Then, for each province and each year, per capita incomes are divided by average of provinces for the corresponding year to generate relative incomes, which are later taken their natural logarithms.

14 This study does not follow stochastic and deterministic convergence definitions based on the deterministic or stochastic trend discussed in Islam (2003) because they may create confusion with the stochastic and deterministic convergence I define here.

15 For an assessment of cross-section and time-series approaches to convergence see Bernard and Durlauf (1995 and 1996).

16 See Bakış (2018) for details.

17 Income per capita was first announced in 2016 and revised in 2020 for 2009-2019. Chain volume index was adopted in 2016.

18 Chen and Nordhaus (2011) grade countries from A to E in terms of output and luminosity compliance where A is the highest grade while E is the lowest. Turkiye has the grade C and luminosity has small value added in A, B, and C due to high measurement error. Therefore, the extended income per capita series by Düşündere (2019 and 2020) may have no significant information additions to the subnational income per capita series. There are strong evidences to support such that for all years and provinces, correlation between the predicted and official income per capita ranges between 99.38% and 99.9% (Düşündere, 2019). Besides, official data in 2020, 2021 and 2022 are not used in this study owing to different sources.

19 This dataset is constructed on behalf of The Economic Policy Research Foundation of Turkiye (TEPAV).

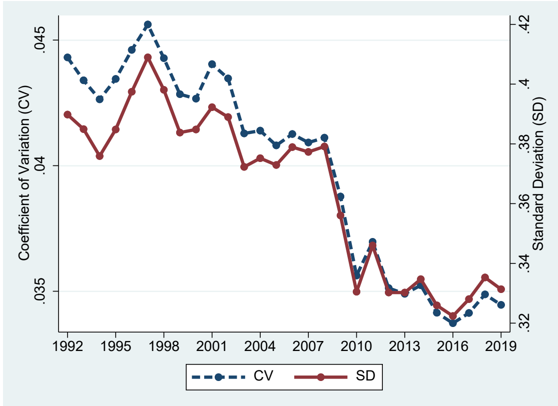

Figure 1: σ- convergence

Figure 1 presents σ-convergence using standard deviation and coefficient of variation. The 1990s were characterized by relatively higher income dispersion among provinces. After 2000, a radical fall in statistics can be seen that is equivalent to an improvement in income distribution. The decline in income dispersion intensified during the 2008-2010 period, which can be attributed to the global financial crisis. Thus, it may signal convergence towards low – income provinces. Indeed, σ-convergence does not tell where the provinces heading to low income or high-income. Figure 2 and 3 represent choropleth maps about average real income levels and real GDP growth rates from 1992 to 2019. East and West distinction is explicitly monitored. Eastern Anatolia and the South-east Anatolia stay at the lowest quartile, whereas Western provinces are located at the highest quartile. There is a smooth transition from high-income to low-income provinces. Tunceli, Erzincan, Trabzon, Rize, and Artvin disturb this smooth transition. Differences among provinces are eroded during that time period in favor of the North-west Anatolia, according to Figure 3. There are some individual units as well in which growth rates belong to the highest quartile and no significant pattern is observed.

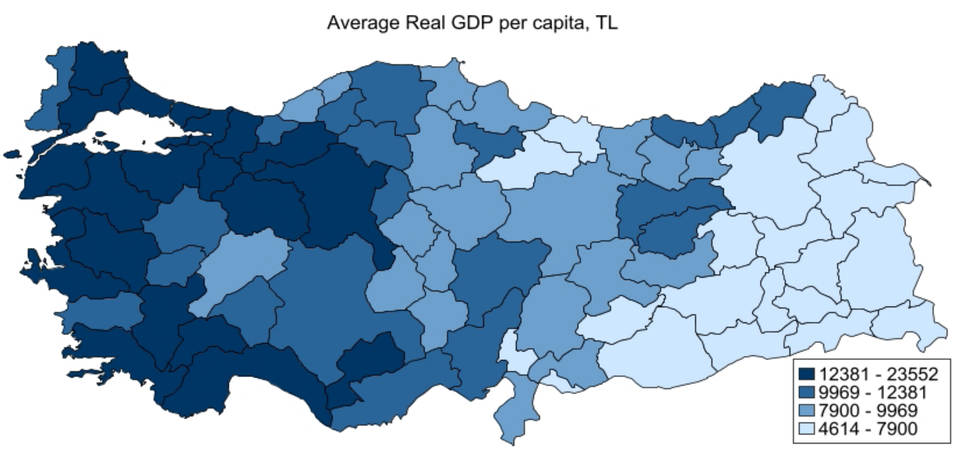

Figure 2: Average Real GDP per capita, TL

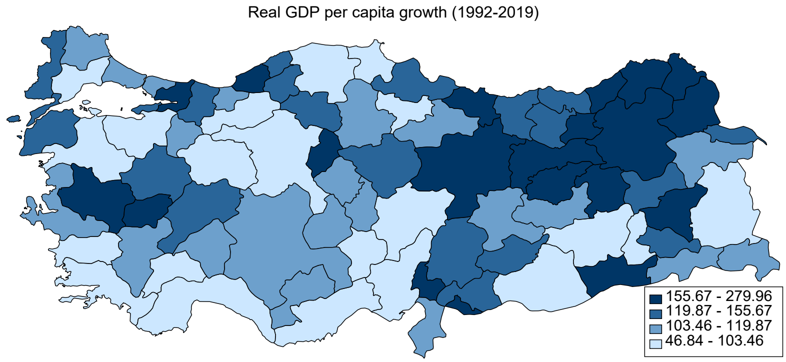

Figure 3: Real GDP per capita Growth Rate (1992-2019)

5. Econometric Methodology

Bai and Ng (2005) underline the importance of the null hypothesis of stationarity, which is more natural than the null hypothesis of a unit root for many economic problems. It can be argued that if convergence is rejected for the stationary null, this would provide stronger evidence against the convergence hypothesis than simply failing to reject the unit root null hypothesis (Bai and Ng, 2005). Becker, Enders, and Lee (2006) also shed more light on this debate by pointing out that tests with the null hypothesis of a unit root have low power in stationarity when a theory has to be tested under the null of stationarity. Therefore, I follow in their footsteps, and a battery of stationarity tests has been implemented to check the regional income convergence dynamics. In addition, instead of merely univariate tests, a common practice in the literature, panel tests that utilize more information are used as the provinces are adjacent to each other and likely to be affected to varying degrees by the same shocks. Besides panel outcomes, a dual perspective is adopted due to the possibility of interpretations of individual series in terms of stationarity. Univariate time series stationarity tests suffer from low power, while panel counterparts can enhance the power due to a higher number of observations but can be difficult to interpret (Maddala, 1999; Smith and Fuertas, 2010). First of all, the information is always obtained from univariate tests; thus, as Maddala (1999) proposed, movement to panel tests may not solve the varying conclusions, but more powerful tests can be a natural remedy. Therefore, this study challenges the recent empirical literature of (stochastic) convergence in Turkiye on the grounds of a series of tests considering potential maladies that can harm the power of the tests.

5.1. No-shift: Hadri (2000) and Cross-Sectional Dependence

Hadri (2000) extends the residual-based Lagrange multiplier (LM) univariate stationary test of Kwiatkowski, Phillips, Schmidt, and Shin (1992) [20] and introduces panel data stationarity test with

20 Smith and Fuertas (2010) emphasize that Kwiatkowski, Phillips, Schmidt and Shin (1992), hereafter, KPSS is sensitive to the bandwidth selection. Unless, it is reported, all bandwidths for spectral window are set to $4(T/100)^{2/9}$.

the null hypothesis of series are stationary around a deterministic trend against the alternative hypothesis of unit root. The model can be written as follows:

$$y_{it} = z_t' \delta_i + r_{it} + \epsilon_{it} \qquad (5)$$

$$r_{it} = r_{i,t-1} + u_{it} \qquad (6)$$

where $\delta_i = [\alpha_i, \beta_i]'$ and $z_t = [1, t]'$ with the trend model. r_{it} is a random walk. $u_{it} \sim IIN(0, \sigma_u^2)$ and $\epsilon_{it} \sim IIN(0, \sigma_\epsilon^2)$ are mutually independent normal, and independent and identically distributed across i and over t. The stationarity null hypothesis is $\sigma_u^2 = 0$ against the alternative of $\sigma_u^2 > 0$. The initial values of r_{i0} are heterogenous fixed unknowns and the trend model can be written as

$$y_{it} = r_{i0} + \beta_i t + \sum_{t=1}^{t} u_{it} + \epsilon_{it} \qquad (7)$$

Partial sum of residuals (S_{it}) is obtained from equation (7) using OLS. The LM test that is the average of the Kwiatkowski *et al.* (1992) test statistic across i, allowing heteroskedasticity, and estimated using the below formula

$$LM = \frac{1}{N}\left(\sum_{i=1}^{N}\left(\frac{1}{T^2}\frac{\sum_{t=1}^{T} S_{it}^2}{\widehat{\sigma_{\epsilon i}^2}}\right)\right) \qquad (8)$$

The benchmark panel test statistic, which is the normalized version of (8), is computed as Z. The above test statistic is normalized to obtain the benchmark panel test statistics

$$Z = \frac{\sqrt{N}(LM - \xi)}{\zeta} \sim N(0,1) \qquad (9)$$

where ξ is the mean and ζ^2 is the variance with 1/15 and 11/6300, respectively (Hadri, 2000). Z is standard normal; thus, there is no need to compute a new set of critical values. Such stationary or unit root tests with presumed cross-sectional independence are first-generation tests. Hadri (2000) panel stationarity test is deliberately preferred in this work because new features are added into the same structure in each stage.

In contrast to spatial economics, where cross-correlation is related to geographic factors such as distance, location, and space, this study treats contemporaneous correlation stemming from unobserved global shocks, local interactions, or pure idiosyncratic correlation among individuals (Moscone and Tosetti, 2009).

The existence of common shocks and unobserved common components pave the way for interdependencies across cross-sectional units (De Hoyos and Sarafidis, 2006). Cross-sectional dependence and potential structural breaks can result in inconsistent and biased inferences. Besides, such issues will also determine what kind of panel unit root or stationary tests have to be adopted. The recently flourishing literature suggests two approaches to identifying cross-sectional dependence (Ditzen, 2021): direct testing for the CD (Pesaran, 2015) and estimating the strength of the dependence (Bailey, Kapetanios, and Pesaran, 2016). Both methods detect the cross-sectional dependence in relative GDP per capita. First, Pesaran (2015, 2021) test statistic is estimated using the following equation:

$$CD = \left[\frac{TN(N-1)}{2}\right]^{1/2} \left(\frac{2}{N(N-1)} \sum_{i=1}^{N-1} \sum_{j=i+1}^{N} \widehat{\rho_{ij}}\right) \quad (10)$$

where $\widehat{\rho_{ij}}$ is the pair-wise correlation coefficient. However, Pesaran (2015) offers to shift the null hypothesis of cross section independence of Pesaran (2004) with weak [21] cross-sectional dependence for panels with large N. The null hypothesis can be shown as $0 \leq \alpha < (2-d)/4$, and α measures the degree of cross-sectional dependence (Pesaran, 2015). In other words, CD test examines for $\alpha < 0.5$[23] (Ditzen, 2021). So, with 81 provinces, I can safely use the null hypothesis of weak cross-sectional dependence against strong cross-sectional dependence. According to Table 1, weak convergence cannot be rejected as the p-value is greater than 0.10. As an alternative way to gauge the cross correlation, CD* test of Pesaran and Xie (2023) which a bias corrected version of Pesaran (2015) is estimated using the following equation:

$$CD^* = \frac{CD + \sqrt{\frac{T}{2}}\theta_n}{1-\theta_n} \quad (11)$$

where θ_n is the bias-corrected term. The result of CD* ends up with strong cross-sectional dependence. Although, outcomes of CD and CD* are enough to justify the utilization of panel tests capturing cross-correlations, as a final attempt to settle the degree of cross-correlation, the exponent of cross-sectional dependence is estimated using Bailey *et al.* (2016), which has quite decent small sample property. This approach tries to determine the value of α from the range of [0,1]. The range of [0.5,1] corresponds to different degrees of strong cross-sectional dependence, while the range of [0,0.5] corresponds to different degrees of weak cross-sectional dependence. It would be more appropriate to verify the degree of cross-sectional dependence is sufficiently large, that is to say, $\alpha > 1/2$, to justify the use of Bailey *et al.* (2016) method. Here CD^* test can be referred to better interpret $\mathring{\alpha}$ in Table 1. $\mathring{\alpha}$ is the bias-adjusted estimator [22] of α, which is close to 1, implying strong cross-sectional dependence.

Table 1: Testing Cross-Sectional Dependence

CD	CD*	$\mathring{\alpha}_{0.05}$	$\mathring{\alpha}$	$\mathring{\alpha}_{0.95}$
1.01	-1.68	0.851	0.918	0.986
(0.315)	(0.093)		[0.041]	

Notes: Numbers in parentheses are p-values. The number in brackets is the standard error. The first 4 principal components are used in the estimation of CD*. $\mathring{\alpha}_{0.05}$ and $\mathring{\alpha}_{0.95}$ give the 90% confidence interval bands.

21 Weak cross-sectional dependence means that the correlation between units at each point in time converges to zero as the number of cross sections goes to infinity. Under strong dependence the correlation converges to a constant.

22 The details can be found in Bailey *et al.* (2016).

5.2. No-shift: Hadri and Kurozumi (2011)

Hadri and Kurozumi (2011 and 2012) modify the data-generating process of Hadri (2000) and incorporate cross-sectional dependence in the form of a common factor. Error component ϵ_{it} is redefined as following

$$\epsilon_{it} = f_t \gamma_i + v_{it} \tag{12}$$

f_t is a one-dimensional latent common factor, and each individual is very likely to be affected by the common factor with the loading factor γ_i. To eliminate cross-sectional dependence, Pesaran (2007) methodology is followed. Cross-sectional average of the model, composed of (5), (6), and (12), is taken to remove the common factor [23]. New partial sum of residuals (S_{it}^{w}) is constructed from the cross-sectional average model using OLS. Then, using Hadri (2000) procedure, same statistics are obtained as in (8) and (9) but to differentiate the notation, *A* subscripts are added as LM_A and Z_A. Individual test statistics are seen in the innermost parenthesis in (8), and that term is divided by a consistent long-run variance estimator to correct for serial correlation, so the innermost parenthesis is replaced by $\frac{1}{\sigma_i^2 T^2} \sum_{t=1}^{T} S_{it}^{w^2}$. As suggested in Hadri and Kurozumi (2012), that estimator is chosen following Sul, Phillips, and Choi (2005) to enhance the power of the test, especially for the trend case. This study applies Sul *et al.* (2005) with quadratic spectral specification.

Beyond cross-sectional dependence, another problem potentially undermining the power of the test, is well documented in Perron (1989) and Lee, Huang, and Shin (1997), may arise due to erroneous omission of structural breaks. Lee *et al.* (1997) depict that stationarity tests ignoring the potential structural break(s) are biased towards rejecting the stationarity null hypothesis and create a size distortion problem. Alongside this, mis-specified placing and numbering of the breaks can severely distort the power of the test; thus, to refrain from such complications, a stationarity test, Carrion-i-Silvestre, *Barrio-Castro, and Lopez-Bazo*. (2005), that can endogenously determine both number and location of breaks. This test also addresses cross-sectional dependence through the nonparametric bootstrapping of Maddala and Wu (1999).

5.3. Sharp-shift: Carrion-i-Silvestre, Barrio-Castro, and Lopez-Bazo (2005)

Carrion-i-Silvestre *et al.* (2005) attach two new components to the random walk process of equation (2) in the form of dummy variables as the changes in the level and slope to capture the date of the break(s). Equations (5) and (6) are adjusted in line with $\delta_i = [\alpha_i, \beta_i]'$ and $z_t = \left[1, t, DU_{i,1,t}, \ldots, DU_{i,m_i,t}, DT_{i,1,t}^{*}, \ldots, DT_{i,m_i,t}^{*}\right]'$. For reasons of parsimony, under the null hypothesis the data generating process of the model with shifts in the mean and time trend is assumed to be

$$y_{it} = r_{it} + \beta_i t + \sum_{s=1}^{m_i} \theta_{i,s} DU_{i,s,t} + \sum_{s=1}^{m_i} \gamma_{i,s} DT_{i,s,t}^{*} + \epsilon_{it} \tag{13}$$

23 In order to save space, averaged model is not added but can be seen in Hadri and Kurozumi (2011 and 2012).

The dummy variable $DT^{*}_{i,s,t} = t - T^{i}_{b,s}$ for $t > T^{i}_{b,s}$ and 0 otherwise, where $s = 1, \dots, m_i$ and m_i is the maximum number of structural breaks imposed, and $T^{i}_{b,s}$ is the sth date of the break for the individual i. The dummy variable $DU_{i,s,t} = 1$ for $t > T^{i}_{b,s}$ and 0 otherwise. The null hypothesis of stationarity is slightly modified compared to Hadri (2000) and Hadri and Kurozumi (2011) to $\sigma^2_{u,i} = 0$ against the nonstationary alternative of $\sigma^2_{u,i} > 0$. Partial sum of residuals is obtained from equation (13) again using OLS. As it is built upon the framework of the Hadri (2000), equation (8) is estimated for $LM(\lambda)$ where λ stands for the dependence of the test on the break dates. Finally, $Z(\lambda)$ is assessed by rewriting $LM(\lambda)$ for LM in equation (8) for the panel test statistics. $Z(\lambda)$ can also be calculated by assuming that long-run variance is homogeneous across individuals. The number of breaks is estimated using LWZ criterion as suggested by Carrion-i-Silvestre *et al.* (2005) when trending regressors are included. Long run variance estimator in our analysis is Sul, Phillips, and Choi (2005) with quadratic spectral quadratic spectral and m is set to 5.

5.4. Smooth-shift: Nazlıoğlu and Karul (2017)

Tests directly identifying the number of breaks, location of breaks, or even their functional form examine the phenomenon of *sharp break*s with the help of time dummies. However, such time dummy practices may not be enough to fully comprehend and transmit the true nature of breaks. The trend is considered to consist of sections that are linear between breaks, while discontinuity is in the realm of possibility (Enders and Lee, 2004). Thus, false specifications of breaks can be as detrimental as their total ignorance. As has been a common topic of debate recently (Enders and Lee, 2004; Becker *et al.* 2006), many macroeconomic time series are characterized by rather *smooth break*s or *gradual break*s, corresponding to structural breaks with an unknown number of breaks, dates, duration, and functional form. The Fourier approximation can mimic various forms of structural breaks or nonlinearities in the deterministic term [24] (Becker *et al.*, 2006). Nazlıoğlu and Karul (2017) borrow the univariate framework of Becker *et al.*, (2006), extend it, and build their novel panel stationarity test. They insert a deterministic term as a function of time as $z_i(t)$ instead of $z_t'\delta_i$ into the DGP in equation (5). The model below is slightly different from Becker *et al.* (2006) as it includes the common factor.

$$y_{it} = z_i(t) + r_{it} + f_t\gamma_i + v_{it} \tag{14}$$

A Fourier expansion with a single frequency component, as in equation (15), is capable of constructing a level and trend shift model.

$$z_i(t) = \alpha_i + \beta_i t + \gamma_{1i}\sin\left(\frac{2\pi kt}{T}\right) + \gamma_{2i}\cos\left(\frac{2\pi kt}{T}\right) \tag{15}$$

k is the Fourier frequency component, and $r_{i0} = 0$ for all i. γ_{1i} measures the amplitude and displacement of shifts is captured by γ_{2i}. As opposed to sharp breaks, smooth breaks using the

24 A strictly linear trend is just a special case.

Fourier approximation has a weakness arising out of unknown form, numbers, and dates of breaks is that it is not possible to analyze the changes of the values of the constant and time trend before and after the structural changes (Tsong, Lee, Tsai and Hu, 2016), which has a vast empirical literature on it starting with Tomljanovich and Vogelsang (2002). Individual test statistics are computed using the following equation

$$\tau_{\tau_i}(k) = \frac{1}{T^2}\frac{\sum_{t=1}^{T}\tilde{S}_{it}(k)^2}{\tilde{\sigma}_{vi}^2} \tag{16}$$

where $\tilde{S}_{it}(k)$ is the sum of OLS residuals from equation (14) and $\tilde{\sigma}_{vi}^2$ is the long run variance [25]. The average of individual statistics (τ_τ) is taken to obtain the below panel test statistic.

$$FP(k) = \frac{1}{N}\left(\sum_{i=1}^{N}\left(\tau_{\tau_i}(k)\right)\right) \tag{17}$$

The null hypothesis of stationarity converges to the standard normal distribution. Thus, the final version of panel test statistic is defined as

$$FZ(k) = \frac{\sqrt{N}(FP(k)-\xi(k))}{\zeta(k)} \sim N(0,1) \tag{18}$$

Values of $\xi(k)$ and $\zeta^2(k)$ for constant, and constant and trend models can be found in Table 1 in Nazlioglu and Karul (2017). The long-run variance is estimated with the Bartlett kernel with Kurozumi (2002), as suggested by Nazlioglu and Karul (2017), due to their superior performance over the rule of Sul et al. (2005).

Which frequency has to be preferred needs great attention and depends upon the sort of data. As argued by Becker, Enders, and Hurn (2004), highly persistent macroeconomic data requires the value of k as 1 or 2 to control for breaks [26] and test for the stationarity versus non-stationarity, where the higher frequencies are not associated with structural breaks but stochastic parameter variability. Nazlioglu and Karul (2017) assume homogenous [27] frequency across cross-sections in order to obtain the asymptotic distribution of panel statistics. According to Lee, Wu, and Yang (2016) homogenous frequency does not mean identical breaks across cross-sections.

25 For the details of the long run variance see Becker *et al*. (2006).

26 "It is difficult to distinguish between a structural break and certain types of nonlinearities. Clearly, a series with a break can be viewed as a special case of a process that is nonlinear in its parameters. As such, our approach can be viewed as an attempt to provide a general procedure to approximate unknown nonlinear components (Becker *et al*. 2006: p.2)"

27 Proper frequency selection especially in time series is possible through grid-search by minimizing sum of squared residuals (Becker *et al*. 2006). To the best my knowledge, similar procedure is not available for panel case.

6. Results

According to Hadri (2000) test statistics, stochastic convergence is not observed in 38 provinces, while shocks have only a transitory impact on 41 provinces across the country but are mostly located in the Mediterranean and Eastern Black Sea regions. There is no clear-cut East-West distinction in terms of convergence. The number of provinces converging to the country average slightly increases to 45 provinces. Although the outcome of 32 provinces does not change when cross-sectional dependence is controlled, the bias that may arise due to erroneous omission of this facet is eliminated. The discrepancy between stochastically convergent provinces according to no-shift models is quite obvious. However, the novelty of this study is the merging of information obtained from univariate test statistics and panel test statistics simultaneously. Panel B parts of Tables 2, 3, and 4 depict panel tests. Panel A parts of Tables 2, 3, and 4 reflect univariate test statistics for each province. Yet, univariate tests in Panel A indicate that only concentrating on the panel level may hide the inner dynamics; thus, this also validates the approach adopted in the study. The null hypothesis of stationarity is rejected in Table 2, leading to, to some extent, divergence at the panel level.

On the other hand, the time span is 28 years, which is quite long for a developing country such as Turkiye. Many significant economic crises (1994 and 2001) stemmed from inner sources or (2007-2008) transmitted from the world during that period may disrupt the estimation of tests that ignore structural breaks. To avoid this and mitigate potential issues, it is necessary to capture the underlying dynamics by allowing the test to account for structural breaks. The groundbreaking feature of Carrion-i Silvestre et al. (2005) is the endogenous determination of structural breaks, and the restriction in front of the number of breaks is removed. Table 3 shows that incorporating structural breaks improved the number of provinces with stochastic convergence to 54. Once again, metropolitan cities such as Istanbul, Ankara, İzmir, and Bursa appear as consistently convergent irrespective of the panel unit root test. In 13 provinces, there is no structural break and most structural breaks take place in Gümüşhane with 5 breaks. Dates of structural breaks vary across the country, but mostly, they correspond to economic crisis periods. The effect of the 1994 crisis may be detected in 1995 and later years; the 2000-2001 crisis is less visible as a break, but as a caveat, it should be noted that 1999 is the catastrophic earthquake year and subsequent years can capture this. Besides, the following years can experience this demolition on economy as a prolonged shock. 2007-08 global financial crises also spilled over and appeared as a break for some provinces. Local, national, and presidential election years should be monitored carefully as a potential source of breaks in this respect.

The difference in structural break dates may also signal out that the same shock may have different impacts on the regions. Panel tests are estimated for two separate cases in Table 3, cross-sectional independence and cross-sectional dependence. Assuming cross-sectional independence results in no stochastic convergence overall, and this conclusion is robust to both homogenous and heterogenous long-run variances. To take into consideration the cross-sectional dependence, bootstrap critical values are obtained, and according to Panel C, homogeneity in the long-run variance ends up with a rejection of stationarity while heterogeneity leads to stationarity in the panel. Therefore, strong interpretations are not possible here.

Smooth-shift models are displayed in Table 4 for k stands for Fourier frequencies. Following Becker, et al. (2004), higher frequencies are not suitable for usage. Once again, at the panel level, FzK statistics reject the stochastic convergence. On the other hand, for 24 provinces, individual test statistics are in the rejection region of the 10% critical value, leading to stochastic convergence. The remaining 54 provinces do not follow a convergence path. When the panel stationarity test is conducted at k=2, the number of convergent provinces falls to only 18. As a result, when cross-sectional dependence is controlled and instead of sharp shift, smooth shifts are allowed in the model, stochastic convergence at the provincial level weakens significantly. Additionally, findings from univariate cases approach to the panel findings, which are robust to the model selection. Unlike the country basis analysis of which outcomes of the panel tests are sensitive to the selection of the panel members such as missing data, membership of an organization, or interest of researchers in particular countries (Ford, Jackson and Kline, 2006), working with a single country and its regions, to some degree, help us to avoid such a problem. However, this study admits that socio-cultural and socio-economic factors may have great importance in settling this convergence issue.

Table 2: No Shift Model

Panel A: province-by-province tests							
		Hadri (2000)	**Hadri and Kurozumi (2012)**			**Hadri (2000)**	**Hadri and Kurozumi (2012)**
Nuts3	**Province**	**KPSS**	**KPSS**	**Nuts3**	**Province**	**KPSS**	**KPSS**
TR100	İstanbul	0.068	0.072	TR811	Zonguldak	0.070	0.045
TR211	Tekirdağ	0.104	0.101	TR812	Karabük	0.117	**0.126**
TR212	Edirne	**0.182**	**0.172**	TR813	Bartın	0.074	0.088
TR213	Kırklareli	0.071	0.070	TR821	Kastamonu	0.058	0.076
TR221	Balıkesir	**0.136**	0.123	TR822	Çankırı	**0.142**	**0.140**
TR222	Çanakkale	**0.163**	**0.148**	TR823	Sinop	**0.138**	**0.150**
TR310	İzmir	0.076	0.113	TR831	Samsun	**0.128**	**0.135**
TR321	Aydın	0.094	0.104	TR832	Tokat	**0.127**	**0.162**
TR322	Denizli	0.102	0.108	TR833	Çorum	0.060	0.077
TR323	Muğla	**0.147**	**0.132**	TR834	Amasya	0.088	0.092
TR331	Manisa	0.089	**0.147**	TR901	Trabzon	0.077	0.055
TR332	Afyonkarahisar	0.117	0.113	TR902	Ordu	0.073	0.066
TR333	Kütahya	0.112	0.109	TR903	Giresun	**0.182**	**0.180**
TR334	Uşak	0.098	0.091	TR904	Rize	**0.126**	0.098
TR411	Bursa	0.115	**0.151**	TR905	Artvin	**0.162**	**0.156**
TR412	Eskişehir	0.043	0.043	TR906	Gümüşhane	**0.148**	**0.147**
TR413	Bilecik	**0.128**	0.118	TRA11	Erzurum	0.098	0.053
TR421	Kocaeli	0.062	0.092	TRA12	Erzincan	**0.135**	**0.134**
TR422	Sakarya	0.076	0.110	TRA13	Bayburt	0.104	**0.137**
TR423	Düzce	0.092	0.074	TRA21	Ağrı	**0.126**	0.103
TR424	Bolu	0.082	0.092	TRA22	Kars	0.074	0.055
TR425	Yalova	**0.154**	**0.169**	TRA23	Iğdır	0.122	0.095
TR510	Ankara	0.103	0.100	TRA24	Ardahan	0.041	0.093
TR521	Konya	**0.167**	**0.157**	TRB11	Malatya	**0.127**	**0.146**
TR522	Karaman	0.064	0.061	TRB12	Elazığ	0.095	0.078
TR611	Antalya	**0.166**	**0.154**	TRB13	Bingöl	**0.165**	**0.153**
TR612	Isparta	**0.160**	**0.133**	TRB14	Tunceli	**0.173**	**0.168**
TR613	Burdur	**0.173**	**0.160**	TRB21	Van	**0.161**	**0.145**
TR621	Adana	**0.159**	**0.132**	TRB22	Muş	0.086	0.097
TR622	Mersin	**0.154**	**0.154**	TRB23	Bitlis	0.092	0.065
TR631	Hatay	**0.139**	**0.153**	TRB24	Hakkari	**0.189**	**0.174**
TR632	Kahramanmaraş	0.043	0.068	TRC11	Gaziantep	**0.146**	**0.158**
TR633	Osmaniye	**0.171**	**0.155**	TRC12	Adıyaman	**0.165**	**0.160**
TR711	Kırıkkale	0.069	0.071	TRC13	Kilis	**0.127**	0.122
TR712	Aksaray	**0.191**	**0.188**	TRC21	Şanlıurfa	0.085	0.086
TR713	Niğde	**0.153**	**0.155**	TRC22	Diyarbakır	0.083	0.057
TR714	Nevşehir	**0.127**	**0.154**	TRC31	Mardin	**0.169**	**0.150**
TR715	Kırşehir	0.061	0.074	TRC32	Batman	0.082	0.069
TR721	Kayseri	0.105	0.101	TRC33	Şırnak	**0.205**	**0.192**
TR722	Sivas	0.070	0.119	TRC34	Siirt	**0.134**	0.103
TR723	Yozgat	0.063	0.096				

Panel B: panel tests		
	Stat.	**p-value**
	10.634	0.000
Hadri (2000)	**10.691**	0.000
Hadri and Kurozumi (2011)		

Notes: The critical values are 0.119, 0.146, and 0.216 for 10%, 5%, and 1%, respectively. The bold numbers show the rejection of the null hypothesis of stationarity.

Table 3: Sharp Shift Model

Panel A: province-by-province tests											
		Carrion-i Silvestre *et al.* (2005)							**Bootstrap Critical Values**		
Nuts3	**Province**	**KPSS**	**m**	$T_{b,1}$	$T_{b,2}$	$T_{b,3}$	$T_{b,4}$	$T_{b,5}$	**0.90**	**0.95**	**0.99**
TR100	İstanbul	0.063	2	2000	2009				0.076	0.173	0.331
TR211	Tekirdağ	0.093	2	1999	2009	2010	2010		0.586	0.774	1.173
TR212	Edirne	0.157	3	1999	2005	2015	2015		0.162	0.180	0.218
TR213	Kırklareli	0.036	2	1999	2008	2014	2014		0.299	0.435	0.665
TR221	Balıkesir	**0.406**	3	1999	2008	2006			0.125	0.153	0.313
TR222	Çanakkale	0.107	2	1996	2012	2010			0.170	0.209	0.358
TR310	İzmir	0.065	0	1997	2010	2009			0.262	0.379	0.724
TR321	Aydın	0.036	2	1996	2008	2013			0.308	0.375	0.519
TR322	Denizli	0.037	2	1996	2007	2006			0.190	0.259	0.368
TR323	Muğla	0.076	1	2009	2001	2007			0.354	0.434	0.572
TR331	Manisa	**0.561**	1	1999	2008	2007			0.116	0.164	0.362
TR332	Afyonkarahisar	0.041	1	2007	2009				0.154	0.193	0.323
TR333	Kütahya	0.043	1	1996	2013				0.112	0.132	0.250
TR334	Uşak	0.070	3	1995	2007				0.167	0.207	0.265
TR411	Bursa	0.085	4	2005	2009				0.563	0.615	0.725
TR412	Eskişehir	0.171	0	1998	2012				0.260	0.355	0.601
TR413	Bilecik	0.069	1	2005	2013				0.159	0.180	0.240
TR421	Kocaeli	**0.158**	2	2013	2015				0.126	0.263	0.503
TR422	Sakarya	0.038	2	1996	2007				0.103	0.169	0.284
TR423	Düzce	0.073	1	1997	2005				0.219	0.327	0.564
TR424	Bolu	**0.257**	2	1999	2001				0.193	0.229	0.311
TR425	Yalova	0.043	2	1999	2008				0.135	0.182	0.267
TR510	Ankara	0.072	2	2009	2007				0.073	0.108	0.275
TR521	Konya	**0.764**	2	1996	2009				0.160	0.181	0.299
TR522	Karaman	0.050	2	2000	2002				0.086	0.103	0.182
TR611	Antalya	0.064	2	1996	2011				0.189	0.219	0.382
TR612	Isparta	0.055	1	2003	2014				0.245	0.311	0.463
TR613	Burdur	0.092	2	2000	2002				0.198	0.252	0.360
TR621	Adana	**0.469**	1	1999	2002				0.148	0.194	0.342
TR622	Mersin	**0.218**	1	1997					0.143	0.164	0.264
TR631	Hatay	**0.216**	1	1995					0.168	0.218	0.356
TR632	Kahramanmaraş	**0.402**	3	1999					0.078	0.128	0.165
TR633	Osmaniye	**0.388**	4	1998					0.126	0.144	0.172
TR711	Kırıkkale	0.052	1	1997					0.239	0.331	0.518
TR712	Aksaray	0.086	3	1997					0.166	0.177	0.208
TR713	Niğde	0.049	1	1997					0.280	0.334	0.480
TR714	Nevşehir	0.211	0	1999					0.252	0.346	0.597
TR715	Kırşehir	0.118	0	1995					0.247	0.352	0.593
TR721	Kayseri	0.081	2	1999					0.217	0.283	0.406
TR722	Sivas	0.093	2	2004					0.141	0.279	0.500
TR723	Yozgat	0.175	1	2002					0.184	0.275	0.421
TR811	Zonguldak	0.105	0	1995					0.261	0.356	0.598
TR812	Karabük	0.184	0	1995					0.266	0.368	0.628
TR813	Bartın	**0.225**	0						0.265	0.383	0.641
TR821	Kastamonu	0.162	3						0.399	0.452	0.578
TR822	Çankırı	0.100	1						0.172	0.200	0.324
TR823	Sinop	0.082	2						0.132	0.224	0.424
TR831	Samsun	**0.170**	2						0.081	0.116	0.195
TR832	Tokat	**1.698**	4						0.074	0.098	0.138
TR833	Çorum	0.174	3						0.418	0.485	0.635
TR834	Amasya	0.146	0						0.245	0.330	0.557

Notes: The bold numbers show the rejection of the null hypothesis of stationarity. The number of breaks is selected using the LWZ criteria. Bootstrap critical values are obtained with 4000 replications.

Table 3: Sharp Shift Model (continued)

Panel A: province-by-province tests											
		Carrion-i Silvestre *et al.* (2005)							**Bootstrap Critical Values**		
Nuts3	**Province**	**KPSS**	**m**	$T_{b,1}$	$T_{b,2}$	$T_{b,3}$	$T_{b,4}$	$T_{b,5}$	**0.90**	**0.95**	**0.99**
TR901	Trabzon	0.104	0						0.249	0.351	0.589
TR902	Ordu	**0.170**	2	2004	2014	2005	2011	2015	0.100	0.119	0.188
TR903	Giresun	**0.296**	1	2004	1999	2011	2015		0.146	0.165	0.212
TR904	Rize	0.087	0	2010	2005	2008			0.250	0.357	0.623
TR905	Artvin	**0.298**	1	1995	2001	2012			0.159	0.177	0.248
TR906	Gümüşhane	**0.962**	2	1995	2015	2014			0.523	0.572	0.661
TRA11	Erzurum	0.040	1	1999	2008	2013			0.369	0.428	0.544
TRA12	Erzincan	**0.917**	4	1995	2011				0.247	0.325	0.582
TRA13	Bayburt	0.097	0	1995	2009				0.246	0.357	0.643
TRA21	Ağrı	0.041	1	2002	2012				0.332	0.397	0.552
TRA22	Kars	0.038	1	1996	2010				0.372	0.431	0.552
TRA23	Iğdır	0.046	1	2005	2007				0.184	0.246	0.395
TRA24	Ardahan	0.028	0	1999	2012				0.259	0.366	0.637
TRB11	Malatya	0.099	0	2000	2009				0.255	0.355	0.635
TRB12	Elazığ	**0.085**	3	1995	2015				0.073	0.081	0.107
TRB13	Bingöl	**0.271**	2	2000	2009				0.228	0.326	0.527
TRB14	Tunceli	0.086	2	2000	2009				0.138	0.173	0.351
TRB21	Van	0.168	2	1995	2008				0.411	0.517	0.708
TRB22	Muş	0.035	2	2000					0.204	0.236	0.302
TRB23	Bitlis	0.165	2	2010					0.190	0.248	0.394
TRB24	Hakkari	**0.613**	2	2002					0.113	0.167	0.325
TRC11	Gaziantep	**0.429**	3	2007					0.414	0.468	0.567
TRC12	Adıyaman	**0.183**	2	1999					0.101	0.110	0.160
TRC13	Kilis	0.051	1	2000					0.164	0.187	0.251
TRC21	Şanlıurfa	**0.291**	2	2008					0.087	0.115	0.216
TRC22	Diyarbakır	**0.625**	2	2000					0.375	0.492	0.719
TRC31	Mardin	**0.353**	2						0.108	0.120	0.207
TRC32	Batman	0.044	3						0.111	0.170	0.267
TRC33	Şırnak	0.144	1						0.180	0.233	0.361
TRC34	Siirt	**0.222**	3						0.069	0.111	0.395
Panel B: panel tests assuming cross-section independence											
		Stat.	**p-value**								
LM(λ)-homogenous		**70.187**	0.000								
LM(λ)-heterogenous		**19.163**	0.000								
Panel C: panel tests assuming cross-section dependence (bootstrap distribution)											
				0.90	**0.95**		**0.99**				
LM(λ)-homogenous				25.016	26.076		28.101				
LM(λ)-heterogenous				42.749	44.874		49.616				

Notes: The bold numbers show the rejection of the null hypothesis of stationarity. The number of breaks is selected using the LWZ criteria. Bootstrap critical values are obtained with 4000 replications.

Table 4: Smooth Shift Model

Panel A: province-by-province tests							
		Nazlioglu and Karul (2017)				**Nazlioglu and Karul (2017)**	
		k=1	**k=2**			**k=1**	**k=2**
Nuts3	**Province**	**KPSS**	**KPSS**	**Nuts3**	**Province**	**KPSS**	**KPSS**
TR100	İstanbul	**0.072**	0.092	TR811	Zonguldak	0.036	0.038
TR211	Tekirdağ	**0.059**	**0.126**	TR812	Karabük	0.044	**0.126**
TR212	Edirne	**0.074**	**0.147**	TR813	Bartın	**0.056**	**0.142**
TR213	Kırklareli	**0.072**	0.092	TR821	Kastamonu	**0.058**	**0.110**
TR221	Balıkesir	**0.054**	**0.145**	TR822	Çankırı	**0.052**	**0.145**
TR222	Çanakkale	**0.049**	**0.152**	TR823	Sinop	**0.063**	**0.142**
TR310	İzmir	0.040	**0.134**	TR831	Samsun	0.040	**0.142**
TR321	Aydın	**0.051**	**0.114**	TR832	Tokat	0.042	**0.142**
TR322	Denizli	0.038	**0.121**	TR833	Çorum	**0.063**	**0.106**
TR323	Muğla	**0.082**	**0.170**	TR834	Amasya	0.036	0.066
TR331	Manisa	0.035	**0.147**	TR901	Trabzon	**0.062**	0.080
TR332	Afyonkarahisar	0.044	**0.142**	TR902	Ordu	**0.053**	0.080
TR333	Kütahya	0.042	**0.132**	TR903	Giresun	0.036	**0.191**
TR334	Uşak	**0.069**	**0.153**	TR904	Rize	0.041	0.100
TR411	Bursa	**0.065**	**0.143**	TR905	Artvin	**0.063**	**0.148**
TR412	Eskişehir	**0.065**	0.055	TR906	Gümüşhane	**0.059**	**0.139**
TR413	Bilecik	0.029	**0.129**	TRA11	Erzurum	**0.063**	0.033
TR421	Kocaeli	**0.062**	**0.107**	TRA12	Erzincan	**0.059**	**0.139**
TR422	Sakarya	**0.056**	**0.129**	TRA13	Bayburt	**0.052**	**0.170**
TR423	Düzce	**0.048**	0.096	TRA21	Ağrı	**0.064**	**0.135**
TR424	Bolu	**0.052**	**0.109**	TRA22	Kars	**0.055**	0.038
TR425	Yalova	0.031	**0.146**	TRA23	Iğdır	**0.062**	**0.159**
TR510	Ankara	**0.081**	**0.140**	TRA24	Ardahan	0.034	0.091
TR521	Konya	**0.055**	**0.163**	TRB11	Malatya	0.044	**0.180**
TR522	Karaman	**0.054**	0.036	TRB12	Elazığ	**0.055**	**0.120**
TR611	Antalya	**0.056**	**0.158**	TRB13	Bingöl	**0.048**	**0.136**
TR612	Isparta	**0.065**	**0.168**	TRB14	Tunceli	**0.054**	**0.169**
TR613	Burdur	**0.077**	**0.151**	TRB21	Van	0.043	**0.169**
TR621	Adana	**0.055**	**0.140**	TRB22	Muş	**0.064**	**0.107**
TR622	Mersin	**0.049**	**0.139**	TRB23	Bitlis	**0.079**	**0.119**
TR631	Hatay	0.035	**0.190**	TRB24	Hakkari	**0.071**	**0.141**
TR632	Kahramanmaraş	0.035	0.055	TRC11	Gaziantep	**0.068**	**0.141**
TR633	Osmaniye	**0.056**	**0.147**	TRC12	Adıyaman	**0.057**	**0.142**
TR711	Kırıkkale	**0.065**	0.080	TRC13	Kilis	0.040	**0.138**
TR712	Aksaray	**0.048**	**0.156**	TRC21	Şanlıurfa	0.038	0.064
TR713	Niğde	**0.061**	**0.222**	TRC22	Diyarbakır	**0.052**	0.058
TR714	Nevşehir	0.036	**0.174**	TRC31	Mardin	**0.062**	**0.143**
TR715	Kırşehir	0.039	0.074	TRC32	Batman	**0.057**	**0.114**
TR721	Kayseri	**0.083**	**0.143**	TRC33	Şırnak	0.024	**0.145**
TR722	Sivas	**0.051**	**0.132**	TRC34	Siirt	**0.074**	**0.164**
TR723	Yozgat	**0.059**	**0.133**				
Panel B: panel tests							
		Stat.	**p-value**				
FzK (k=1)		**16.912**	0.000				
FzK (k=2)		**17.381**	0.000				

Notes: The bold numbers show the rejection of the null hypothesis of stationarity. k represents the Fourier frequency. Critical values are 0.0471 (10%), 0.0546 (5%), and 0.0716 (1%) for k=1; 0.1034, 0.1321, and 0.2022 for k=2.

7. Conclusion

Turkiye shows up as a very suitable candidate for analyzing the convergence phenomenon due to flagrant regional disparities manifested in the West and East. Despite the voluminous literature, there is no consensus on the presence or type of convergence. This research adheres to the findings of the aforementioned empirical literature on Turkiye, and partially documents some supportive outcomes using the choropleth maps and σ-convergence. Economic growth spreads unevenly at the provincial level, making catch-up challenging for initially low-income areas. For example, the Northeastern part overperformed compared to the rest in terms of real per capita growth. On the one hand, generally, low-income areas experienced relatively low progress. On the other hand, income dispersion gets narrower, especially after the 2000s. The speed of σ-convergence soared up around 2008. As a matter of fact, this can provide some evidence for approaching the level of lower-income regions as a country.

This study follows an alternative formulation and perspective to shed more light on convergence in Turkiye between 1992 and 2019. It adopts a more statistical attitude and adds new flavors to estimation mechanics in pursuit of the best data-generating process of income per capita. As a result, this study focuses on stochastic convergence. However, some limitations arise in this approach as other economic factors and initial income levels cannot be controlled, but dealing with the pure data itself and its relative ratio to the province averages contributes to understanding of convergence from a different angle. Incorporating data and region-specific elements like cross-sectional dependence, endogenously determined structural breaks, and smooth breaks requires implementing a set of panel stationarity tests. Four different panel stationarity tests are employed, constructed on the same structure. Therefore, no methodological probable inconsistency exists, and results can be directly comparable. The panel framework permits the study to examine the univariate cases as well. Thus, empirical results are interpreted in two layers. The results of the panel stationarity tests partially track the literature, and Turkiye, with its provinces, does not have stochastic convergence. Controlling different potential features of the data also does not alter that conclusion. However, stochastic convergence or stochastic divergence is not omnipresent at the provincial level, and there is no regional pattern. Provinces should be discussed in-depth to reveal the reasoning behind this absence. There is also a tendency towards obtaining fewer provinces as stochastically convergent. The results, in particular, from univariate cases, demand a lot of care, and perhaps further technical carving out.

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