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Financial Stress and the Trade Balance in Turkey: Empirical Evidence from Credit and Real Exchange Rate Channels

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Abstract

In this study, the effects of fluctuations and crises in financial markets on the trade balance were analyzed using monthly data from the period 2000–2025. In the model established, the trade balance was treated as the dependent variable, while the Financial Pressure (Stress) Index was used as the primary independent variable to represent stress and fragility in financial markets. In order to explain the mechanisms through which financial stress affects the trade balance, domestic credit and the real effective exchange rate index were included in the model as transmission channels. Empirical findings reveal the existence of a stable equilibrium relationship between the variables in the long term, according to the Maki cointegration test, which allows for multiple structural breaks. The cointegration coefficients obtained show that global uncertainties and financial crises negatively affect the trade balance under structural breaks and cause fluctuations in trade performance. However, the VAR-based impulse-response functions indicate that the impact of external shocks on financial stress and other variables dissipates in the short term, within approximately 2–3 months. On the other hand, Vector Error Correction Model (VECM) results estimated based on the VAR model show that financial stress, domestic credit, and exchange rate channels have significant long-term effects on the trade balance. These findings reveal that the trade balance is sensitive not only to short-term shocks but also to the overall health and stability of the financial system. In this context, it is concluded that multidimensional and comprehensive policy measures are needed to limit the negative effects of global financial uncertainty and crises on foreign trade, which strengthen financial stability, direct the credit mechanism in favor of production and exports, and aim to reduce exchange rate volatility.

Keywords: *Financial Stress, Trade Balance, Credit Channel, Real Effective Exchange Rate, Time Series Econometrics*

Türkiye'de Finansal Stres ve Ticaret Dengesi: Kredi ve Reel Döviz Kuru Kanallarından Elde Edilen Ampirik Kanıtlar
Öz

Bu çalışmada, finansal piyasalardaki dalgalanmalar ve krizlerin dış ticaret dengesi üzerindeki etkileri, 2000–2025 dönemi aylık verileriyle analiz edilmiştir. Modelde dış ticaret dengesi bağımlı değişken, finansal piyasalardaki stres ve kırılganlıkları temsilen Finansal Baskı (Stres) Endeksi temel bağımsız değişken olarak kullanılmıştır. Finansal baskının dış ticaret dengesi üzerindeki etkilerini açıklayabilmek amacıyla, yurt içi krediler ve reel efektif döviz kuru endeksi modele aktarım kanalları olarak dâhil edilmiştir. Ampirik bulgular, çoklu yapısal kırılmalara izin veren "Maki" eşbütünleşme testine göre değişkenler arasında uzun dönemde istikrarlı bir denge ilişkisi vardır. Eşbütünleşme katsayıları, küresel belirsizlikler ve finansal krizlerin yapısal kırılmalar altında dış dengeyi olumsuz etkilediğini ve performansında dalgalanmalara yol açtığını göstermektedir. VAR modeline dayalı etki–tepki fonksiyonları, değişkenlerde meydana gelen dışsal şokların etkisinin kısa vadede, yaklaşık 2–3 aylık bir süre içerisinde sönümlendiğine işaret etmektedir. Öte yandan, VAR modeli temelinde tahmin edilen Vektör Hata Düzeltme Modeli (VECM) sonuçları, finansal baskı, yurt içi krediler ve döviz kuru kanallarının dış ticaret dengesi üzerinde uzun dönemde anlamlı etkiler yarattığını göstermektedir. Bu bulgular, dış ticaret dengesinin yalnızca kısa vadeli şoklara değil, finansal sistemin genel sağlığına ve istikrarına duyarlı olduğunu ortaya koymaktadır. Bu çerçevede, küresel finansal belirsizliklerin ve krizlerin dış ticaret üzerindeki olumsuz etkilerini sınırlayabilmek için finansal istikrarı güçlendiren, kredi mekanizmasını üretim ve ihracat lehine yönlendiren ve döviz kuru oynaklığını azaltmayı hedefleyen çok boyutlu ve kapsayıcı politika tedbirlerine ihtiyaç duyulduğu sonucuna ulaşılmaktadır.

Anahtar Kelimeler: *Finansal Stres, Dış Ticaret Dengesi, Yurt İçi Krediler, Reel Efektif Döviz Kuru, Zaman Serisi Ekonometrisi*

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1. Introduction

The financial system plays a critical role in the functioning of the real economy. It serves as a core mechanism that channels savings into investment, facilitates risk sharing, and regulates the allocation of resources among economic agents. In recent decades, its influence on the real sector has become more visible. Key macroeconomic variables such as production, employment, and foreign trade have grown increasingly sensitive to developments and expectations in financial markets. Today, finance and economic growth can almost be described as closely linked. Finance helps reduce frictions in the real sector that arise from transaction and information costs. By doing so, it supports the savings and investment decisions of economic actors and, in turn, contributes to capital accumulation and economic growth (Levine, 2005). This interaction operates through several channels, including the credit channel, asset prices, expectations, and risk premia. Changes in financial conditions directly influence firms' decisions regarding investment, production, and trade. Especially in an environment of deepening global financial integration, fluctuations within the financial system can generate rapid and often simultaneous effects on the real economy (Reinhart and Rogoff, 2009).

The proper functioning of the financial system is essential for the efficient allocation of resources, the formation of reasonable financing costs, and the preservation of trust among economic agents. A well-functioning financial system enables firms to make production and investment decisions within a predictable financing environment, while also guiding households' saving and consumption decisions in a stable manner. In this context, financial resilience refers to the capacity of the financial system to withstand internal and external shocks, absorb these shocks, and continue performing its core functions without major disruption (Thakor, 2014). By contrast, disruptions in the financial system can generate significant costs for the real economy through credit contraction, rising uncertainty, higher risk premia, and deteriorating expectations. When the financial intermediation mechanism weakens, firms face difficulties in accessing finance, investments are postponed, and production processes are interrupted. Such a situation of financial fragility does not only harm growth and employment; it also negatively affects trade flows and overall macroeconomic stability, thereby deepening existing economic vulnerabilities (UNDP, 2024).

Given the decisive role of the financial system in the real economy and the importance of financial resilience, it becomes necessary to systematically monitor deteriorations in financial conditions and the accumulation of stress. Imbalances in financial markets rarely emerge through a single indicator. Instead, they tend to build up through multiple channels such as interest rates, risk premia, asset prices, and market volatility, and eventually spill over to the real economy. For this reason, measuring pressures and tensions within the financial system in a comprehensive framework is crucial for assessing financial stability and analyzing possible macroeconomic effects (Ilgın, 2024). This need has led to the development of indices in the literature that capture financial stress in a multidimensional way. Efforts to gather information about the overall condition or health of the financial sector through various indicators date back to the early 1980s. Early studies generally focused on a single market or a single indicator. However, in the 1990s, there was a growing need for measures capable of capturing total stress across the entire financial system. Since the late 1990s, the number of such studies has increased, many of them conducted by central banks in advanced economies and international financial institutions (Öztürkler and Göksel, 2013).

With technological progress, the speed of financial flows has increased, and both financial intermediation activities and financial instruments have changed considerably. This transformation has made the monitoring and regulation of financial markets more complex. At the same time, as financialization has expanded beyond the size of the real economy, the interaction between the financial system and the real sector has become deeper and more intertwined. As a result, feedback mechanisms between these two spheres have strengthened. However, identifying the exact nature of these relationships and the transmission channels through which they operate has become more challenging (Yetiz and Ünal, 2018). Following the 2007–2009 global financial crisis in particular, financial stability has emerged as a central research area for both policymakers and academics. In this context, measuring financial stress and developing stress indices have gained importance for central banks aiming to monitor financial risks. Nevertheless, because financial stress has a multidimensional and somewhat abstract structure, measurement efforts in this field remain limited, and many of the proposed indices are largely experimental in nature. Despite these limitations, such indices are expected to provide valuable insights into financial conditions, especially in the design and implementation of macroprudential policies (Çamlıca and Güneş, 2016).

These indices, often referred to as “financial stress” or “financial pressure” measures, are constructed through the integration of various sub-market indicators. The Financial Stress Index aims to capture tensions in financial markets by combining different financial variables—such as interest rates, risk premia, asset prices, and market volatility—into a single composite indicator. Pressures in financial markets and expectations of potential losses are considered the main sources of financial stress (Illing and Liu, 2003). Rising financial stress, together with increasing risk and uncertainty, usually reflects a fragile financial structure and the emergence of shocks. When the index takes positive values, this indicates growing pressure and stress in financial markets, higher risk perceptions, and tighter financial conditions. In contrast, negative values point to relatively loose financial conditions, stronger risk appetite, and lower pressure on the financial system. In this sense, the Financial Stress Index serves as an important leading indicator for assessing the current state of the financial system and its possible effects on the real economy (Hakkio and Keeton, 2009).

The financial crisis that emerged in the United States in the late 2000s and quickly spread across the globe accelerated research focusing on the structure and vulnerabilities of the financial sector. The relative strength of financial markets does not mean that financial stress cannot arise. Both internal and external shocks can significantly increase financial stress, even in economies with well-developed financial systems. Financial stress can act as a trigger for both market volatility and real economic slowdown. In this context, efforts to measure financial stress and to develop and monitor Financial Stress Indices (FPIs) have gained importance. The measurement of financial stress has attracted considerable attention in the literature, particularly after the global financial crisis. Existing studies show that most financial stress indices are calculated for advanced economies—especially the United States—and generally at a monthly frequency. In contrast, FPI applications for emerging economies are relatively new. Given the increasing integration of these countries into the global financial system, the need to monitor financial stress using higher-frequency data has become more evident (Ekinci, 2013).

The impact of financial pressure and stress on foreign trade operates through several transmission channels. First, through the credit channel, rising financial stress reduces banks' risk appetite and makes it more difficult for firms to access export financing. This creates a limiting effect on trade volumes, especially for exporters that rely heavily on external finance and for small and medium-sized enterprises (SMEs). Second, through the uncertainty and expectations channel, higher financial stress leads firms to postpone investment and production decisions. This, in turn, negatively affects the supply of goods and services subject to export. Third, during periods of financial stress, increased exchange rate volatility and higher risk premia raise the costs of trade contracts and make international transactions more risky. Although exchange rate depreciation caused by financial stress may, in some cases, support exports in the short run through improved price competitiveness, the literature generally suggests that this effect is temporary. In the medium and long term, rising uncertainty and higher financing costs tend to exert a restrictive effect on foreign trade. Therefore, the overall impact of financial stress on foreign trade is mostly negative (Kaya and Kılıç, 2017).

In the remainder of the study, the theoretical framework is presented first. This study, which examines the impact of the financial stress index on the trade balance, is important in several respects. It seeks to clarify the transmission mechanisms between financial conditions and the real economy, to analyze the sensitivity of trade performance to financial vulnerabilities, and to provide policy-relevant implications regarding the relationship between financial stability and external balance. The third section reviews the relevant literature and summarizes its main contributions. The fourth section introduces the main model and the data used in the analysis. The fifth and sixth sections explain the econometric methods and present the results of the empirical analysis conducted to assess the validity of the main model. Finally, the findings of the study are discussed in the seventh and concluding section.

2.Theoretical Framework

This section presents the theoretical background of the study. Financial stress refers to a situation in which the financial system experiences disruptions in its core intermediation functions, leading to heightened uncertainty, increased risk premia, liquidity shortages, and tighter credit conditions. Unlike a financial crisis, which denotes a discrete systemic breakdown, financial stress represents a continuum of tension within financial markets that may or may not culminate in a crisis. In this sense, financial stress captures both observable market volatility and latent fragility within the financial structure.

Theoretically, it reflects a deterioration in financial resilience and a weakening of the mechanisms that channel savings into productive investment.

To empirically capture financial stress, this study employs the Financial Pressure Index (FPI), which aggregates standardized movements in exchange rates, interest rates, and foreign reserves. Each of these components corresponds to a specific theoretical transmission channel. Exchange rate volatility reflects external fragility and capital flow reversals; interest rate fluctuations capture tightening liquidity conditions and rising risk premia; and changes in international reserves indicate pressures on the balance of payments and central bank intervention capacity. By combining these indicators into a composite index, the FPI operationalizes the multidimensional nature of financial stress in a measurable and dynamic framework.

Thus, the empirical variable used in the analysis is not an arbitrary proxy but a theoretically grounded composite indicator that reflects the interaction between currency markets, monetary conditions, and external vulnerability. It is widely accepted that pressure and stress in the financial system affect real economic activity and foreign trade performance through several channels, including economic growth, credit conditions, expectations, uncertainty, and exchange rates. In particular, during periods of financial stress, rising risk perceptions and higher financing costs become key factors shaping firms' trade decisions and the path of the trade balance. These indicators also serve as leading signals of economic crises. Market participants are expected to interpret these signals in advance and adjust their financial positions accordingly. This tendency became more pronounced after the crises experienced in the 1990s. Some studies aim to predict crises by using a country's own macroeconomic variables, while others examine multiple countries to identify variables that may significantly explain crisis episodes. In general, these studies have employed Logit–Probit models, the KLR model (Kaminsky, Lizondo, and Reinhart&Aneta 2006), and Artificial Neural Network (ANN) models. First, the financial pressure index is formulated as follows:

$$FBE_t = \frac{\left[\dfrac{e_t - \mu_e}{\sigma_e}\right] + \left[\dfrac{r_t - \mu_r}{\sigma_r}\right] + \left[\dfrac{i_t - \mu_i}{\sigma_i}\right]}{3} \tag{1}$$

where, the Financial Pressure Index (FPI) is defined as an index obtained by taking the average of the standardized percentage changes in the exchange rate (USD), the overnight interest rate, and the Central Bank's gross foreign exchange reserves. When this index exceeds a certain threshold value, it is interpreted as indicating the presence of a financial crisis; otherwise, it is assumed that no crisis has occurred. In the denominator, e_t represents the percentage change in the exchange rate, r_t denotes the percentage change in the Central Bank's gross foreign exchange reserves, and i_t indicates the percentage change in the overnight interest rate. As shown, each variable included in the FPI is standardized by subtracting its arithmetic mean from each observation and dividing the result by its standard deviation. Since the weights of the variables are assumed to be equal, the sum of the three standardized values is divided by three. A crisis is then identified as a period that leads to an increase in the FPI. For this purpose, the threshold value is defined as:

$T.V. = \mu + a\sigma,$

where μ is the arithmetic mean of the FPI, σ is its standard deviation, and a is a constant. Periods in which the FPI exceeds this threshold are classified as crisis periods, and a dummy variable takes the value K = 1. Otherwise, the period is considered non-crisis, and the dummy variable takes the value K = 0 (Sevim, 2012: 56–57). An important point in this framework is the ability to capture early warning signals by detecting deviations from normal patterns in the months preceding a crisis. A signal is generated when the value of any independent variable exceeds its predetermined threshold level.

One of the most influential studies using the signal approach is the work of Kaminsky, Lizondo, and Reinhart&Aneta (2006), commonly referred to as the "KLR Model." In their study, Kaminsky et al. (1998) examined empirical evidence on financial crises and developed an early warning system for crisis prediction. Within this framework, a crisis signal is generated when the pressure index variable exceeds a predetermined threshold value. In addition to financial and monetary variables—such as money supply, short-term capital flows, budget balance, and foreign currency deposit accounts—the model also incorporates several external sector indicators. These include the ratio of imports to GDP, the current account balance, the ratio of net exports to GDP, changes in the terms of trade, changes in exports and imports, and the export-to-import coverage ratio. Another method used in this context is the Artificial Neural Networks (ANN) model. The main difference from the previous approach is its ability to capture non-linear relationships more effectively. Similar to this method, external trade indicators are also employed leading variables in this framework (Franck & Schmied, 2003).

The link between financial stress and the trade balance operates through three interrelated channels. First, the credit channel suggests that heightened financial stress reduces banks' risk appetite, restricts access to working capital and export financing, and suppresses trade volumes. Second, the domestic absorption channel implies that tightening financial conditions compress domestic demand, particularly import demand, thereby mechanically improving the trade balance. Third, the exchange rate channel reflects the impact of currency volatility on price competitiveness and external trade costs. These channels interact dynamically, and their relative strength may vary across short- and long-term horizons. Since the construction of the Financial Pressure Index (FPI) includes not only the real sector and the banking sector but also the foreign exchange market, measurement has also been conducted using the Exchange Market Pressure Index (EMPI).

$$EMPI_{i,t} = \frac{(\Delta e_{it} - \mu_{i,\Delta e})}{\sigma_{i,\Delta e}} - \frac{(\Delta RES_{it} - \mu_{i,\Delta RES})}{\sigma_{i,\Delta RES}} \tag{2}$$

In the equation, $\Delta e_{i,t}$ represents the change in the real exchange rate, while $\Delta RES_{i,t}$ denotes the change in international reserves. The symbol σ refers to the standard deviation of the relevant variable, and μ indicates its mean. As can be seen from the equation, an increase in the real exchange rate or a decrease in international reserves leads to a rise in the Financial Pressure Index (FPI).

In the literature, one of the most commonly used methods for constructing a financial stress index is the equal variance weighting method. Due to its simplicity and ease of implementation, it is widely preferred. Under this weighting scheme, all sub-indices are assigned equal weights, implying that each sub-market is assumed to have equal importance in identifying periods of stress within the overall financial system (Koyunlu, 2019: 32). For the standardization process, the mean of the series is first subtracted from each observation. The result is then divided by the standard deviation according to the following formula:

$$z_t = \frac{(x_t - \bar{x})}{\sigma} \tag{3}$$

where, z_t denotes the standardized and mean-adjusted pressure (stress) indicator, x_t represents the raw stress indicator for the market, $\bar{x}$ is the sample mean, and σ refers to the standard deviation. After this standardization process, all standardized components are aggregated using their arithmetic average according to the following formula (Balakrishnan et al., 2009):

$$VEW = (\sum_{1}^{k} z_k) / k \tag{4}$$

To determine crisis periods, exceeding a certain threshold level of the FBE is considered a crisis;

$$FPI > 2.5\sigma_{FPI} + \mu_{FPI} \tag{5}$$

$$FPI > 2.5\sigma_{FPI} + \mu_{FPI} \tag{6}$$

where, σ_{FPI} represents the sample standard deviation, and μ_{FPI} denotes the sample mean of the financial stress indices calculated under each of the three methods. If the pressure index exceeds these values, it indicates the presence of a crisis; otherwise, it is considered that no crisis has occurred (Gerni et al., 2005).

3. Literature Review

A review of the related literature shows that most studies focus on calculating Financial Pressure Indices (FPIs) for specific country groups—such as advanced or emerging economies—and on examining whether the FPI can predict financial crises, that is, whether it provides early warning signals. Especially after the 2001 global crisis, research on measuring financial stress gained noticeable momentum. The remaining strand of the literature investigates the relationship between the FPI and other macroeconomic variables, including economic growth, international trade, global capital flows, and foreign direct investment.

There is no standardized method for constructing a financial stress index that applies universally to all countries or regions. Each country differs in terms of macroeconomic conditions, the degree of integration with the global economy, and the development level of its financial system. Consequently, various indices have been developed to reflect the specific economic conditions and financial structures of individual countries. In this context, studies such as Illing and Liu (2006), Nelson and Perli (2007), Hakkio and Keeton (2009), Aruoba et al. (2009), Hatzius et al. (2010), Blix Grimaldi (2010), Oet et al. (2011), Hollo et al. (2012), Gupta et al. (2014), Park and Mercado (2014), Freixas et al. (2015), Romer and Romer (2017), and Leaven and Valencia (2018) developed indices for advanced economies, including Canada and the U.S. On the other hand, studies such as Uygur (2001), Cardarelli et al. (2009), Reinhart and Rogoff (2009), Balakrishnan et al. (2009), Kliesen and Smith (2010), Kota and Saqe (2013), Leaven and Valencia (2020), and Groen et al. (2022) have focused primarily on emerging economies. In addition, there are studies aimed at developing financial stress indices specifically for Turkey. These include Çevik et al. (2012), Elekdağ et al. (2012), Ekinci (2013), Aklan et al. (2015), Kara et al. (2015), Çamlıca and Güneş (2016), Kaya and Kılınç (2017), and Eraslan (2017).

Some studies, focusing specifically on financial markets, have examined the impact of financial pressure on stock exchanges and equity prices. Baron et al. (2021), Dan and Li (2022), Fu et al. (2022), Zhang and Li (2022), Armah, Bossman, and Amewu (2023), Günay, Öner, and Aybars (2023), Liang et al. (2023), and Xu et al. (2023) all find that stock markets and individual stock prices are negatively affected in the medium and long term by increases in financial pressure.

Other studies have focused on the impact of rising financial stress on broader macroeconomic variables. In particular, when a crisis evolves into systemic risk—that is, when financial instability becomes widespread—it inevitably affects key macroeconomic indicators such as unemployment, inflation, economic growth, and exports. In this context, De Bandt and Hartmann (2000) argue that an FPI that has evolved into a systemic crisis can drag numerous institutions—from banks to real production units—into a difficult and persistent negative process, even if some failing entities are removed. Das et al. (2018) examined the relationship between the financial stress index and gold and crude oil prices for the period 1993–2017 using a non-parametric quantitative causality analysis. Their results indicate the existence of a bidirectional causality between the FPI and both oil and gold prices.

Das et al. (2019) investigated whether economic policy uncertainty, geopolitical risk, and financial stress affected emerging economies in a similar way during the 1997–2018 period, using a non-parametric causality test. Their results indicate that these shocks significantly influenced the countries' markets, although the magnitude and nature of the effects varied across different markets. Ahir et al. (2023), in a study covering 110 countries and applying the Romer and Romer (2017) approach, analyzed how increases in the Financial Pressure Index (FPI) affected output, whether there was a causal relationship, and whether the effects differed between advanced and emerging economies. Their findings reveal a causal link between the FPI and economic growth, show that financial stress negatively impacts economic activity, and demonstrate that financial crises triggered by rising FPI have more destructive effects in emerging and developing markets compared to advanced economies.

Similarly, Elekdag and Kanlı (2010) analyze the relationship between financial stress and economic activity from the perspective of emerging markets. They examine the effects of both external and domestic financial stress shocks on Turkish economic activity in a comparative context with other emerging economies. Their findings indicate that, whether temporary or persistent, industrial production levels are significantly affected by financial shocks. This underscores the strong impact of financial stress on economic activity, a conclusion further confirmed by the recent global financial crisis.

The Financial Pressure Index (FPI) is a comprehensive measurement tool that consolidates vulnerabilities, stress dynamics, and systemic risk factors within the financial system into a single indicator, serving as an early warning mechanism. In this respect, the FPI is not merely a static index that describes current financial conditions; it is also a dynamic monitoring tool that generates leading signals about potential future economic fluctuations. Accurately and timely interpreting the signals provided by the index allows for the anticipation of sudden and disruptive effects of economic crises. This, in turn, enables policymakers to design proactive and preventive policy measures. Increases observed in the Financial Pressure Index (FPI) indicate rising risk perceptions in financial markets, higher levels of uncertainty, and the beginning of a weakening in financial intermediation mechanisms. This process is transmitted directly to the real sector through credit channels, placing pressure on investment, production, and trade decisions, and disrupting macroeconomic equilibrium mechanisms.

From the perspective of the trade balance, periods of heightened financial pressure are associated with increased exchange rate volatility, tighter external financing conditions, and more difficult access to trade finance. These factors generate asymmetric effects on export and import volumes and structurally influence the dynamics of the external balance. In this context, the FPI provides a critical analytical framework for examining the network of interactions among the trade balance, credit expansion, exchange rate movements, and economic growth. The significance of this study lies in treating the financial pressure index not merely as an indicator confined to financial markets but as an integrated measure encompassing the real economy, the trade balance, and macroeconomic stability. While accurately interpreting the signals from the FPI cannot fully prevent economic crises, it has strong potential as a policy tool to mitigate the depth of crises, slow their propagation, and limit their economic costs. In this regard, the study aims to highlight the effects of financial pressure dynamics on the trade balance and to provide an analytical and policy-oriented contribution to the literature on early warning systems, macroprudential policy design, and crisis management.

Finally, the study identifies three research questions. First, how do increases in the Financial Pressure Index (FPI) affect Turkey's trade balance in the short and long term? Second, through which transmission channels, namely the real effective exchange rate and domestic credit conditions—do these pressures influence the trade balance? And third, are the effects of shocks arising from financial pressure on the trade balance temporary or persistent, and does the magnitude of these effects change over time in conjunction with structural breaks? These research questions stem from the need to examine financial pressure dynamics not only within financial markets but also in terms of their broader effects on the real economy and the trade balance in a holistic framework. Increases in the Financial Pressure Index (FPI) can influence trade decisions both directly and indirectly through tighter credit conditions, greater exchange rate volatility, and heightened uncertainty. Accordingly, this study investigates how the effects of financial pressure on the trade balance are shaped both directly and through the channels of the real effective exchange rate and domestic credit conditions. In addition, based on the assumption that these effects may differ in the short and long term and can evolve over time due to global crises, financial fluctuations, and structural transformations, the persistence and dynamic nature of financial pressure shocks are also analyzed. In this context, the study aims to fill a significant gap in the literature by empirically testing the role of financial pressure on the trade balance.

4. The Modal and Data

This study examines the impact of financial stress on the trade balance in Turkey. For this purpose, the dependent variable is defined as the trade balance (Trade Balance – TB). The first independent variable is financial stress, represented by the Financial Pressure Index (FPI) calculated for emerging economies. The FPI consists of 33 financial market variables, including yield spreads, valuation measures, and interest rates. When the index equals zero, the variables are at their long-term averages. A positive value indicates rising financial risks, while a negative value suggests that the market is relatively stable and secure. The second independent variable is domestic credit to the private sector (Domestic Credit-DC), used to capture the trade finance channel. Finally, to control for price competitiveness, the real effective exchange rate (REER) is included as the third independent variable in the model.

Table 1. Definition, Construction, and Sources of Variables

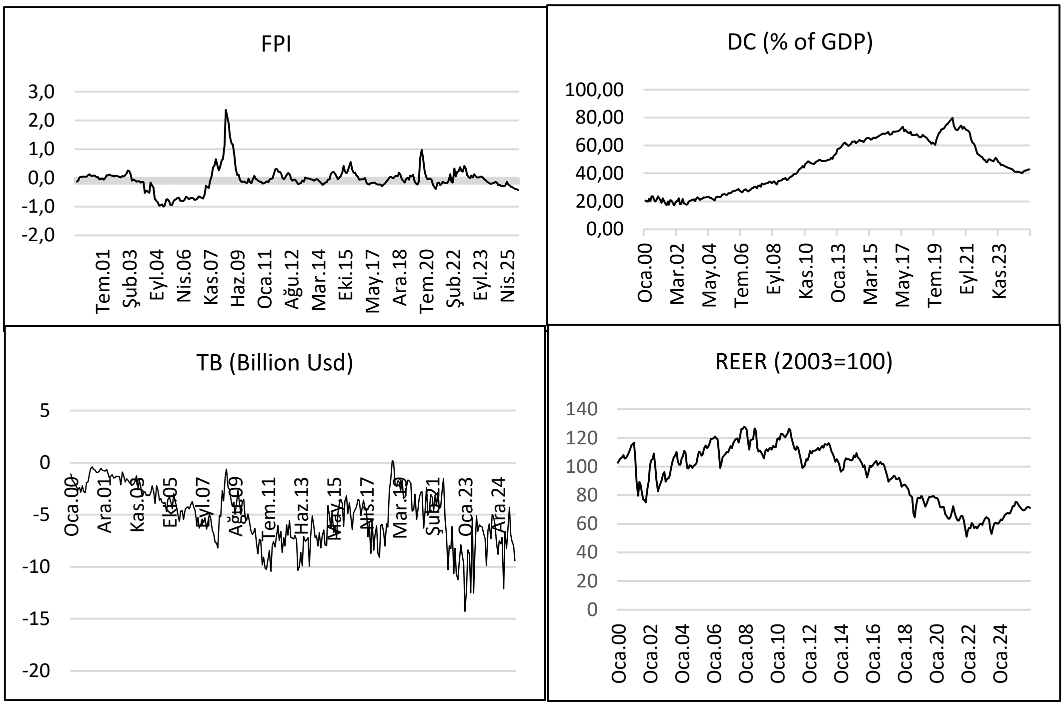
Variable	Definition	Transformation	Construction	Source
TB (Trade Balance)	Difference between exports and imports	Level	TB = Exports − Imports	TURKSTAT / CBRT
FPI (Financial Stress Index)	Composite index capturing financial market stress	Standardized	Constructed from 33 financial market variables including yield spreads, interest rates, and valuation measures	IMF, BIS, National Sources
REER (Real Effective Exchange Rate)	Weighted average of domestic currency against trade partners' currencies, adjusted for inflation	Log / Level	Trade-weighted nominal exchange rate adjusted for relative CPI	BIS, IMF
DC (Domestic Credit)	Credit provided to the private sector by domestic banks	Log / Growth	Domestic credit to private sector (total)	World Bank, CBRT

Accordingly we can express the main model as follows;

$$LnTB_t = \beta_0 + \beta_1 LnFSI_t + \beta_2 LnDC_t + \beta_3 LnREER_t + \varepsilon_t \tag{7}$$

As can be seen, in the model, β_0 represents the constant term, ε denotes the stochastic error term, and the subscript t refers to the monthly time period from January 2000 (2000M1) to December 2025 (2025M12). In this study, the analysis begins in the post-2000 period, which covers a time when global financial and economic uncertainties increased significantly. During this period, events such as the 2001 U.S. recession and the September 11 attacks, the Second Gulf War, the SARS outbreak, the 2007–2008 global financial crisis, the European debt crisis, the Brexit process, the U.S.–China trade wars, the COVID-19 pandemic, and subsequent events—including the global inflation shock, simultaneous monetary tightening by central banks, the Russia–Ukraine war, and vulnerabilities in the banking sector—led to persistent and elevated levels of financial stress.

Figure 1. Time Series Plots of Variables



At first, examining the Financial Pressure Index (FPI) reveals that it reached historically high levels during the 2008–2009 global financial crisis, reflecting a pronounced intensification of financial stress. Although stress levels eased somewhat after the crisis, the index continued to fluctuate around its long-term average. In 2020, financial stress sharply increased again with the onset of the COVID-19 pandemic, though this spike was shorter-lived compared to the 2008 crisis. Following the pandemic, financial stress did not fully dissipate, continuing to exhibit a volatile pattern due to the combined effects of the global inflation shock, monetary tightening, and geopolitical developments.

Second, when examining the ratio of domestic credit to GDP, a clear upward trend is observed from the early 2000s, with financial deepening accelerating particularly after 2010. However, after 2021, there is a noticeable decline in credit ratios, which can be linked to the global monetary tightening process, rising financial uncertainties, and tightening credit conditions. This decline suggests that financial stress may constrain real economic activity and trade through the trade finance channel. In this context, credit dynamics emerge as an important transmission mechanism explaining the effects of financial stress on trade. The third figure represents the trade balance. Observing its trajectory, the balance shows an increasing trend from the mid-2000s, and although it experiences occasional fluctuations, it generally remains at high levels. During global crisis periods and times of intensified financial uncertainty, the trade balance deepens further. Even after 2020, despite rising financial stress, exchange rate volatility, and tighter financing conditions, the trade balance continues to be a persistent issue. This pattern underscores the significant role of financial stress and credit conditions in determining Turkey's trade balance.

Our final figure shows the real effective exchange rate (REER) index. The 2010–2013 period was marked by ample liquidity due to globally accommodative monetary policies, accelerated capital inflows into emerging markets, and a consequent real appreciation of the Turkish lira. This environment stimulated domestic demand and imports in Turkey, contributing to an expansion of the trade balance. However, the REER index has exhibited a pronounced and persistent downward trend since the early 2010s. This reflects a real depreciation of the Turkish lira and indicates that, particularly after 2013, domestic inflation dynamics and exchange rate volatility became key determinants. Despite the theoretical expectation that a real depreciation should improve the trade balance, the persistence of Turkey's trade balance can be explained by the production structure's dependence on imported inputs and the high degree of exchange rate pass-through. This suggests that the relationship between the real exchange rate and the trade balance has weakened due to structural vulnerabilities.

5. Econometric Methods

In this section, the cointegration equation is first estimated to examine whether a long-term equilibrium relationship exists among the variables. For this purpose, a cointegration test is applied based on the time series model defined in equation (1). Structural breaks are also taken into account in this equation. Structural breaks are quite common in time series analyses, often arising from economic shocks that create significant uncertainty. It is well known that analyses conducted using traditional methods, such as Engle and Granger (1987) or Johansen (1991), without considering structural breaks, may produce unreliable results.

In this context, several tests that account for structural breaks have been developed, including those by Gregory and Hansen (1996a), Bai and Perron (1998), and Hatemi-J (2008). However, when the researcher does not have prior information regarding the number of breaks, a test capable of producing reliable results under such conditions is necessary. Therefore, in this study, the residual-based Maki cointegration test (Maki, 2012) is preferred, as it allows for the examination of long-term relationships when the number of structural breaks is unknown.

In the econometric literature, the concept of a "structural" model was first defined by Hurwicz (1962). According to this definition, a model is considered structural if it allows for the prediction of the effects of deliberate policy interventions or changes occurring in the economy or nature. To make such predictions, the model must be able to explain how these interventions generate changes in the model's components—parameters, equations, and observable or unobservable variables. In this context, the Maki test operates under the assumption that the number of unknown breaks in the cointegration vector is less than or equal to a pre-specified maximum number of breaks. The regression models defined for different levels are as follows:

$$y_t = \mu + \sum_{t=1}^{k} \mu_i D_{i,t} + \beta' x_t + u_t \tag{8}$$

$$y_t = \mu + \sum_{t=1}^{k} \mu_i D_{i,t} + \beta' x_t + \sum_{i=1}^{k} B_i' x_t D_{i,t} + u_t \tag{9}$$

$$y_t = \mu + \sum_{t=1}^{k} \mu_i D_{i,t} + \gamma t + \beta' x_t + \sum_{i=1}^{k} B_i' x_t D_{i,t} + u_t \tag{10}$$

$$y_t = \mu + \sum_{t=1}^{k} \mu_i D_{i,t} + \gamma t + \sum_{i=1}^{k} \gamma_i t D_{i,t} \beta' x_t + \sum_{i=1}^{k} B_i' x_t D_{i,t} + u_t \tag{11}$$

According to these equations, $t = 1, 2, \ldots, T$ represents the number of observations. In the models, y_t is the scalar dependent variable and x_t is an m×1 vector of independent variables ($x_{1t}, \ldots, x_{mt}$), both of which are first-order integrated (I(1)), while u_t denotes the error term. μ, μ_i, γ, γ_i, $\beta' = (\beta_1 \ldots \beta_m)$ *and* $\beta_i' = (\beta_{i1} \ldots \beta_{im})$ represent the parameters of the models. Since the Maki test accounts for structural breaks, T_B denotes the break date, k indicates the number of structural breaks, and Drepresents the dummy variable ($t > T_{Bi}$ *(i=1, ⋯, k) ise 1, diğer türlü sıfır)* capturing the structural break.

Within this framework, Equation (8) represents a structural break in the intercept, Equation (9) allows for breaks in both the intercept and the regime (slope coefficients), and Equation (10) captures a level shift with a deterministic trend. Finally, Equation (11) specifies the most comprehensive model, in which breaks in the level, regime, and trend are jointly incorporated. To test the cointegration relationship under the assumption of i breaks (where i ≤ k), the test statistic is first computed (Çınar and Hushmat, 2021).

$$y_t = \mu + \sum_{i=1}^{k} \mu_i D_{i,t} + \beta' x_t + u_t \tag{12}$$

Subsequently, the residuals of the regression are obtained via the Ordinary Least Squares (OLS) method, and an Augmented Dickey–Fuller (ADF) unit root test is applied to these residuals. In this framework, the null hypothesis states that there is no cointegration (i.e., the residuals contain a unit root), expressed as $\rho = 0$, whereas the alternative hypothesis asserts that the residuals are stationary (i.e., no unit root), expressed as $\rho<0$. The model used to estimate these coefficients is specified as follows:

$$\Delta\tilde{u}_t = \rho\tilde{u}_{t-1} + \sum_{j=1}^{p} \propto_j \Delta\tilde{u}_{t-1} + \varepsilon_t \tag{13}$$

For all possible potential break points, the existence of a single structural break is tested under the assumption of p = 0, and the corresponding t-statistic is computed for each candidate break date. Among all observations, the break date that yields the minimum t-statistic (i.e., k = 1) is determined by identifying the observation that minimizes the residual sum of squares in the following equation. This point is defined as the first structural break:

$$SSR_1 = \sum_{t=1}^{T}(y_t - \hat{\mu} - \hat{\mu}_1 D_{1,t} - \hat{\beta}' x_t)^2 \tag{14}$$

In Equation (13), the first structural break date is formally defined as $\hat{b}p_1 = \mathrm{argmin}_{\tau_1^\alpha} SSR_1$.

In order to identify the second possible break from the sub-samples, the following regression equation and its corresponding error term are employed (Bakkal, 2021):

$$y_t = \mu + \hat{\mu}_1 D_{1,t} + \hat{\mu}_2 D_{2,t} + \beta' x_t + u_t \tag{15}$$

$$\Delta\tilde{u}_t = \rho\tilde{u}_{t-1} + \sum_{j=1}^{p} \propto_j \Delta\tilde{u}_{t-1} + \varepsilon_t \tag{16}$$

Subsequently, the T_2^a sub-sample and the τ_2 statistic corresponding to the error term parameter is computed. The second structural break point, bp_2 , is determined by minimizing the sum of squared residuals (SSR_2) over the T_2^a sub-sample.

$$SSR_2 = \sum_{t=1}^{T}(y_t - \hat{\mu} - \hat{\mu}_1 D_{1,t} - \hat{\mu}_2 D_{2,t} - \hat{\beta}' x_t)^2 \tag{17}$$

According to Equation (17), the second structural break point is denoted by $\hat{b}p_2 = \mathrm{argmin}_{\tau_2^\alpha} SSR_2$Here, the conditions for bp_1 and bp_2 that is, the first and second structural break points are applied to all sub-samples in order to determine the corresponding break dates. Subsequently, the procedure defined in Equations (11), (12), and (13) is implemented iteratively until a total of k structural break points are estimated.

In the second stage, a VAR model is estimated in order to derive the Impulse Response Functions (IRFs). If the variables are cointegrated, they move together in the long run, and unexpected shocks to one variable may transmit to others. To analyze such dynamic interactions, impulse response functions are employed. These functions illustrate how uncertainty shocks affect the relevant variables within an autoregressive framework. However, the ordering and structural specification of variables in a standard VAR model may generate substantial bias in the estimation of impulse responses. Moreover, the error terms in VAR systems are typically contemporaneously correlated. In addition, from a policy perspective, issues related to exogeneity assumptions and the graphical interpretation of impulse responses may reduce the reliability of inferences. The results obtained from conventional VAR models are also sensitive to model specification choices, including the number of included variables, deterministic components (such as trends), lag length selection, and data frequency. To address these limitations, Structural VAR (SVAR) models have been developed. This approach imposes theory-based restrictions on the VAR coefficients, allowing the identification of economically meaningful structural shocks. The SVAR framework not only facilitates the examination of causal relationships among variables but also enables the analysis of the effects of isolated structural shocks on the system. Unlike conventional VAR models, the estimated parameters in an SVAR model can be interpreted in a manner consistent with economic theory and policy analysis (Sims, 2002). In this study, the standard VAR model can be structurally represented as follows (Schweikert, 2018):

$$\begin{aligned} WUI_t &= \emptyset_{11} + \emptyset_{12}WUI_{t-1} + \emptyset_{13}CPI_{t-1} + \emptyset_{14}EXP_{t-1} + \varepsilon_{1t} \\ CPI_t &= \emptyset_{21} + \emptyset_{22}WUI_{t-1} + \emptyset_{23}CPI_{t-1} + \emptyset_{24}EXP_{t-1} + \varepsilon_{2t} \\ EXP_t &= \emptyset_{31} + \emptyset_{32}WUI_{t-1} + \emptyset_{33}CPI_{t-1} + \emptyset_{34}EXP_{t-1} + \varepsilon_{3t} \end{aligned} \tag{18}$$

In equation system (15), the error terms are assumed to follow a white noise process with constant variance. Under this assumption, the standard VAR model can be rewritten in structural form by expressing the endogenous variables as functions of their own lagged values and the contemporaneous relationships among them, as follows:

$$\begin{aligned} WUI_t &= \emptyset_{11} + \emptyset_{12}CPI_t + \emptyset_{13}EXP_t + \emptyset_{14}WUI_{t-1} + \emptyset_{15}CPI_{t-1} + \emptyset_{16}EXP_{t-1} + \varepsilon_{1t} \\ CPI_t &= \emptyset_{21} + \emptyset_{22}WUI_t + \emptyset_{23}EXP_t + \emptyset_{24}WUI_{t-1} + \emptyset_{25}CPI_{t-1} + \emptyset_{26}EXP_{t-1} + \varepsilon_{2t} \\ EXP_t &= \emptyset_{31} + \emptyset_{32}WUI_t + \emptyset_{33}CPI_t + \emptyset_{34}WUI_{t-1} + \emptyset_{35}CPI_{t-1} + \emptyset_{36}EXP_{t-1} + \varepsilon_{3t} \end{aligned} \tag{19}$$

In addition, it is assumed that the error terms ε_{1t}, ε_{2t}, and ε_{3t} are mutually uncorrelated. This assumption enables the separate identification of each independent structural shock. When the system of equations in (19) is rearranged by moving the lagged terms and constants to the right-hand side, it can be expressed in matrix form as follows:

$$\begin{bmatrix} 1 & \emptyset_{12} & \emptyset_{13} \\ \emptyset_{22} & 1 & \emptyset_{23} \\ \emptyset_{32} & \emptyset_{11} & 1 \end{bmatrix} * \begin{bmatrix} WUI_t \\ CPI_t \\ EXP_t \end{bmatrix} = \begin{bmatrix} \emptyset_{11} \\ \emptyset_{12} \\ \emptyset_{13} \end{bmatrix} + \begin{bmatrix} \emptyset_{14} & \emptyset_{15} & \emptyset_{16} \\ \emptyset_{24} & \emptyset_{25} & \emptyset_{26} \\ \emptyset_{34} & \emptyset_{35} & \emptyset_{36} \end{bmatrix} * \begin{bmatrix} WUI_t \\ CPI_t \\ EXP_t \end{bmatrix} + \begin{bmatrix} \delta_{11} & 0 & 0 \\ 0 & \delta_{22} & 0 \\ 0 & 0 & \delta_{33} \end{bmatrix} * \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \\ \varepsilon_{3t} \end{bmatrix} \tag{20}$$

The primary objective of structural VAR (SVAR) estimation is to orthogonalize the error terms in order to conduct impulse–response analysis. The restricted SVAR model defines the relationship among the residuals. These residuals represent unanticipated shocks and structural shocks, both of which are exogenous and mutually uncorrelated. Within this framework, ε_{1t}, ε_{2t}, and ε_{3t} are defined as unobservable, zero-mean white noise processes that are serially uncorrelated over time and independent of one another. Equation (20) enables the transformation of the structural form presented in Equation (18) into its reduced-form representation in terms of the vector x_t.

$$Ax_t = A_0 + A_1x_{t-1} + \delta\varepsilon_t \tag{21}$$

Within this framework, the coefficient matrix A represents the structural parameters, x_{t-1} denotes the matrix of lagged variables, δ corresponds to the variance–covariance matrix, and ε_t represents the error term. Although Equation (21) can be used to estimate the structural parameters, identification requires imposing restrictions on the coefficient matrix that is, setting certain coefficients equal to zero. Within this matrix structure, in order to isolate and observe the effect of a particular variable on the dependent variable, the coefficients associated with the remaining explanatory variables are restricted to zero. This identification scheme allows the variables to respond contemporaneously to both domestic and external factors (Villaverde and Ramirez, 2010).

In this study, the primary objective is to estimate the effects of financial stress (FPI), real effective exchange rates (REER), and domestic credit to the private sector (DC) on Turkey's trade balance (TB). Within this framework, the coefficient ϕ_{12} captures the contemporaneous effect of financial stress (FPI) on the trade balance (TB), the coefficient ϕ_{13} reflects the contemporaneous impact of domestic credit (DC) on the trade balance, and ϕ_{14} measures the contemporaneous effect of the real effective exchange rate (REER) on the trade balance. Similarly, the coefficient ϕ_{33} reflects the contemporaneous impact of a change in commodity prices on exports. Finally, the effects of structural shocks on the variables are traced graphically through dynamic multipliers. To this end, Impulse Response Functions (IRFs) are employed. The impulse–response analysis to different structural shocks is a widely used method for revealing the dynamic properties of macroeconomic models. In this context, the Structural Moving Average (SMA) representation of the VAR model is adopted.

$$y_t = v + \sum_{j=0}^{\infty} \Theta_j \varepsilon_{t-j} \tag{22}$$

where Θ_j is independent variable matrix and ε denotes an independently and identically distributed (i.i.d.) error term. If we express this in matrix form, it can be written as follows:;

$$\begin{bmatrix} y_{1t+s} \\ y_{2t+s} \\ y_{3t+s} \end{bmatrix} = \begin{bmatrix} v_1 \\ v_2 \\ v_3 \end{bmatrix} + \begin{bmatrix} \emptyset_{11,0} & \emptyset_{12,0} & \emptyset_{13,0} \\ \emptyset_{21,0} & \emptyset_{22,0} & \emptyset_{23,0} \\ \emptyset_{31,0} & \emptyset_{32,0} & \emptyset_{33,0} \end{bmatrix} \begin{bmatrix} \varepsilon_{1t+s} \\ \varepsilon_{2t+s} \\ \varepsilon_{3t+s} \end{bmatrix} + \cdots + \begin{bmatrix} \emptyset_{11,s} & \emptyset_{11,s} & \emptyset_{11,s} \\ \emptyset_{21,s} & \emptyset_{22,s} & \emptyset_{23,s} \\ \emptyset_{31,s} & \emptyset_{32,s} & \emptyset_{33,s0} \end{bmatrix} \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \\ \varepsilon_{3t} \end{bmatrix} \tag{23}$$

Using these matrices, the structural dynamic multipliers (impulse–response functions) can be obtained in the same manner;

$$\frac{\partial y_{1t+s}}{\partial \varepsilon_{1t}} = \emptyset_{11,s} \;\; ; \;\; \frac{\partial y_{1t+s}}{\partial \varepsilon_{2t}} = \emptyset_{12,s} \;\; ; \;\; \frac{\partial y_{2t+s}}{\partial \varepsilon_{1t}} = \emptyset_{21,s} \;\; ; \;\; \frac{\partial y_{2t+s}}{\partial \varepsilon_{2t}} = \emptyset_{22,s} \tag{24}$$

As can be seen from Equation (24), the coefficients φ_{ij} provide the graphs of the impulse–response functions (IRFs) for each j. In general, IRFs measure the dynamic responses of a given variable to exogenous or unexpected shocks, while holding all other variables in the system constant. In this way, it becomes possible to analyze how one-unit structural shocks that occur in period t affect the other variables throughout the sample period (Björnland, 2000: 9).

In the final stage of the analysis, the causal relationship between the variables and the direction of this relationship are examined using the frequency domain approach. The main reason for preferring this method is that it offers several important advantages compared to the time domain approach. While the time domain approach treats time series as a function of time, the frequency domain analysis evaluates the series as a function of frequency. Hosoya (1991) and Breitung and Candelon (2006) argue that conventional time domain causality tests are insufficient to detect causal relationships that are valid at different frequencies. Causality dynamics may generate different responses across frequencies, and standard Granger causality tests cannot capture these differences. These tests assume that the causality relationship holds across the entire frequency distribution based on a single Wald statistic. However, the causality relationship may differ between the short and long run; a relationship that exists in the short run may disappear in the long run (Görüş and Aydın, 2019). For this reason, instead of time domain tests, the frequency domain approach provides a more comprehensive perspective on the direction and strength of causality at different frequencies. To account for different frequencies, Hosoya (1991) and Yao and Hosoya (2000) introduced a Wald-type causality test based on spectral density into the literature. For this purpose, $z_t = [x_t; y_t]'$ defined a two-dimensional time series vector with observations for t = 1, ..., T. Here, z_t represents a standard VAR model with a finite number of lags (p).

$$z_t = \propto_1 z_{t-1} + \propto_2 z_{t-2} + \cdots + \propto_p z_{t-p} + \varepsilon_t \tag{25}$$

The error term is moved to the other side of the equation and the delay operator (L) is used;

$$\varepsilon_t = \propto_1 z_{t-1} + \propto_2 z_{t-2} + \cdots + \propto_p z_{t-p} - z_t \tag{26}$$

$$\varepsilon_t = z_t(I - \alpha_1 L^1 - \alpha_1 L^2 - . \dots - \alpha_p L^p) \tag{27}$$

where $\theta = 1 - \alpha_1 L^1 - \alpha_2 L^2 - \dots . \alpha_p L^p$ represent polynomial function and $\varepsilon_t = \theta(L) z_t$ shows error term. With $L^k z_t = z_{t-k}$ lag order, we can express 2x2 dimension autoregressive polynomial function as: $\theta = 1 - \alpha_1 L^1 - \alpha_2 L^2 - \dots . \alpha_p L^p$. Here, the error term ε_t is assumed to be a white noise disturbance term, such that $E(\varepsilon_t) = 0$ and $E(\varepsilon_t : \varepsilon'_t) = \Sigma$. Where Σ is positive definite. Using the Cholesky decomposition, a lower triangular matrix G is defined such that $GG' = \Sigma^{-1}$. Its expected value are $E(\eta_t : \eta'_t) = I$ and $\eta_t = G\varepsilon_t$. If the process is stationary, its moving average representation can be written as follows:

$$z_t = \Phi(L)\varepsilon_t = \begin{bmatrix} \Phi_{11}(L) & \Phi_{12}(L) \\ \Phi_{21}(L) & \Phi_{22}(L) \end{bmatrix} \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{bmatrix} \tag{28}$$

$$\psi(L)\eta_t = \begin{bmatrix} \Psi_{11}(L) & \Psi_{12}(L) \\ \Psi_{21}(L) & \Psi_{22}(L) \end{bmatrix} \begin{bmatrix} \eta_{1t} \\ \eta_{2t} \end{bmatrix} \tag{29}$$

Where $\Phi(L) = \theta(L)^1$ and $\Psi(L) = \Phi(L) G^{-1}$. Accordingly, the spectral density function of x_t can be written as follows;

$$f_x(w) = \frac{1}{2\pi} \{ |\psi_{11}(e^{iw})|^2 + |\psi_{21}(e^{iw})|^2 \} \tag{30}$$

In this equation, ω denotes the frequency of the spectral density. Based on this functional relationship, Geweke (1982) and Hosoya (1991) proposed that the causality relationship can be measured as follows:

$$My \to x(w) = \left[log \frac{2\pi f x(w)}{\psi_{11}(e^{-iw})^2} \right] = log \left[log \frac{\psi_{12}(e^{-iw})^2}{\psi_{11}(e^{-iw})^2} \right] \tag{31}$$

If $|\Psi_{12}(e^{-iw})^2 = 0$ or equivalently *log (1)=0* then at frequency ω, variable y does not Granger-cause variable x. If z_t is integrated of order one and the variables are cointegrated, then $\theta(L)$ contains a unit root. If we substract z_t from both side of the equation $\varepsilon_t = \theta(L) z_t$, we get;

$$\Delta z_t = (\theta_1 - I) z_{t-1} + \theta_2 z_{t-2} + \cdots + \theta_p z_{t-p} + \varepsilon_t = \tilde{\theta}(L) z_{t-1} + \varepsilon_t \tag{32}$$

In this equation, $\theta(L) = \theta_1 - I + \theta_2 L^1 + ... + \theta_p L^p$ represent polynomial autoregressive function. When $\theta(L)$ or $\tilde{\theta}(L)$ equal zero, y does not granger cause x.

In this procedure, the null hypothesis of $H_0 : M_{y\rightarrow x}(\omega) = 0$ meaning that there is no causal relationship from x to y at frequency ω is tested against the alternative hypothesis $H_1 : M_{y\rightarrow x}(\omega) \neq 0$ (Yılmaz, 2025).

6. Empirical Findings

In this section of the study, the econometric estimation results are presented. The analysis begins with descriptive statistics of the time series used in the model. First, the mean, median, maximum, and minimum values of the variables are reported. Among the variables, foreign trade exhibits the highest volatility, while the financial stress index shows the lowest volatility. In this study, which examines the impact of the financial pressure index on Turkey's foreign trade, the index is observed to follow a slightly negative average path close to zero over the sample period, reaching a maximum of 2.36 and a minimum of –1.01. Second, the coefficient of variation (CV = σ / μ), that is, the ratio of the standard deviation to the mean (0.41 / 0.06 = 6.8), is higher for the financial stress index compared to the other variables. The relatively high coefficient of variation (approximately 6.8) suggests that the financial stress index does not fluctuate steadily around a stable mean; rather, it displays sharp movements driven by periodic shocks and crisis episodes. However, since the mean of the index is close to zero, this ratio should be interpreted with caution. Third, skewness, kurtosis, and the Jarque–Bera statistics were calculated to evaluate the normality properties of the dataset. Skewness indicates whether the distribution of a series is symmetric, while kurtosis reflects the thickness or thinness of the tails relative to a normal distribution. The Jarque–Bera test jointly assesses skewness and kurtosis to determine whether a series deviates from normality. Failure to reject the null hypothesis in the Jarque–Bera test suggests that the series is normally distributed, whereas rejection indicates a deviation from normality.

According to the test statistics, all variables exhibit negative skewness. However, since the skewness values are close to zero, the distributions can be considered largely symmetric. In addition, the kurtosis values of all variables are positive and close to three, indicating that their tail structures are broadly consistent with a normal distribution. Based on the Jarque–Bera test results, the null hypothesis of normality cannot be rejected at the 5% significance level for all variables except the exchange rate variable. Finally, the correlation matrix is reported. The negative correlation between financial stress and the trade balance (-0.67) indicates that periods of rising financial pressure are associated with a contraction in external imbalances. This finding reflects a mechanism driven more by a decline in imports than by an improvement in export competitiveness. The positive relationship between domestic credit to the private sector and the trade balance (0.57) suggests that credit expansion enlarges the external balance through higher domestic demand and increased imports. Similarly, the positive correlation between the real effective exchange rate and the trade balance implies that real appreciation weakens price competitiveness, encourages imports, and adversely affects the trade balance.

Table 2. Summary Statistics

Stat.	TB	FPI	DC	REER	Correlation Matrix				
Mean	-5070156	-0.06	45.99	94.68		TB	FPI	DC	REER
Median	-4930794	-0.061	46.92	101.14	TB	1	-0.67	0.57	0.42
Max.	187981	2.36	79.82	127.71	FPI	-0.67	1	0.16	-0.66
Min.	-1.4E+07	-1.01	17.08	50.86	DC	0.57	0.16	1	-0.39
St. Dev.	2748391	0.41	18.59	20.21	REER	0.42	-0.66	-0.39	1
Skeweness	-0.08	-0.05	-0.04	-0.02	Covariance Matrix				
Prob.	0.06	0.08	0.11	0.00	FPI	-8861.57	0.1717	1.306	-0.544
Total	1.58E+09	-21.7	14350.12	29540.28	DC	-1.9E+07	1.306	344.54	-146.91

Table 3. Unit Root Tests

Part 1		Level			First Diff.			
Test	DTD	FPI	DC	REER	TB	FPI	DC	REER
ADF	-2.48 (0.12)	-4.48 (0.03)	-1.36 (0.59)	-1.47 (0.08)	-5.19 (0.00)	-13.86 (0.00)	-16.31 (0.00)	-12.96 (0.00)
PP	-4.49 (0.04)	-5.53 (0.07)	-1.89 (0.22)	-1.74 (0.09)	-17.43 (0.00)	-13.49 (0.00)	-16.87 (0.00)	-12.28 (0.00)
DF-GLS	-1.02	-1.83	-1.09	-2.11	-11.98	-8.04	-3.47	-4.89
Part 2								
Level	TB		FPI		DC		REER	
Test	Test İst.	K.D. (5%)	Test İst.	K.D. (5%)	Test İst.	K.D. (5%)	Test İst.	K.D. (5%)
PT	17.54	5.46	9.67	6.05	4.95	4.11	4.22	3.89
MPT	14.14	5.46	8.09	6.05	4.88	4.11	4.04	3.89
MZA	-18.22	-26.15	-18.42	-19.37	-10.08	-25.46	-16.49	-21.63
MSB	0.17	0.16	0.14	0.13	0.16	0.14	0.18	0.12
MZT	-3.25	-3.60	-3.01	-3.81	-3.07	-3.51	-3.02	-3.20
First Diff.	TB		FPI		DC		REER	
Test	Test Stat.	C.V. (5%)	Test Stat.	C.V. (5%)	Test Stat.	C.V. (5%)	Test Stat.	C.V. (5%)
PT	3.51	5.46	5.52	6.05	3.87	4.11	3.18	3.89
MPT	4.05	5.46	5.04	6.05	3.69	4.11	3.09	3.89
MZA	-27.20	-26.15	-32.31	-29.33	-30.08	-25.46	-26.45	-21.63
MSB	0.13	0.16	0.19	0.13	0.11	0.14	0.09	0.12
MZT	-3.68	-3.60	-3.96	-3.81	-3.87	-3.51	-3.44	-3.20
Break Dates	2008M08 /2018M11/2023M02		2003M3 / 2008M11		2020M10		2001M01 / 2010M09 / 2021M12	

Note: In the first section, the probability values associated with the rejection of the null hypothesis (unit root) are reported in parentheses. The critical values for the DF-GLS test are −2.58, −1.94, and −1.61 at the 1%, 5%, and 10% significance levels, respectively. In the multiple structural break test, the maximum lag length used to estimate the long-run variance is set to four, given the monthly frequency of the data.

In the first stage of the analysis, the stationarity properties of the series are examined through unit root tests in order to avoid spurious regression results. In this context, conventional unit root tests—Augmented Dickey–Fuller (ADF) (Dickey and Fuller, 1981), Phillips–Perron (PP) (Phillips and Perron, 1988), and DF-GLS (Elliott et al., 1996)—are employed. In addition, the multiple structural break unit root test developed by Carrion-i-Silvestre et al. (2009) is applied to account for potential structural shifts in the series. The presence of structural breaks may cause standard unit root tests to yield biased results in favor of failing to reject the null hypothesis of a unit root. Accordingly, the results of all applied tests are reported in Table 3.

According to the conventional unit root test results reported in the first section, DTD and REER are stationary at their first differences, whereas FBE and REER appear to be stationary in levels. Specifically, FBE and REER are found to be stationary at the 5% and 10% significance levels in the ADF test, at the 10% level in the PP test, and at the 5% and 10% levels in the DF-GLS test, respectively. However, the results presented in the second section—where structural breaks at the level are taken into account—indicate that all series are integrated of order one, I(1). In other words, the variables are non-stationary in levels but become stationary after first differencing. This finding suggests that when structural breaks are incorporated into the analysis, the Financial Stress Index (FPI) and the exchange rate variable are no longer stationary in levels. The presence of real shifts and significant unexpected shocks during the sample period (with three structural break dates identified for FBE, as reported in the table) explains why these variables appear stationary under conventional tests but are in fact non-stationary once structural breaks are properly considered. An examination of the time series plots indicates that the variables exhibit different structural break patterns over the sample period. First, the Financial Stress Index (FPI) shows a sharp spike during 2008–2009, corresponding to the global financial crisis, and a second notable break around 2020, associated with the COVID-19 pandemic. This pattern suggests that financial stress is highly sensitive to major global shocks. The domestic credit-to-GDP ratio displays a long upward trend, followed by a clear shift in direction after 2020, with one dominant structural break. This change appears consistent with the impact of post-pandemic monetary and macroprudential policies on credit dynamics. The real effective exchange rate (REER) index presents a more complex pattern. It shows multiple structural break tendencies—approximately three—linked to the post-2001 crisis recovery period, the 2008 global financial crisis, and the exchange rate adjustments observed after 2018. This confirms that the REER is responsive to both global financial conditions and domestic macroeconomic and inflationary dynamics. Overall, the visual evidence from the time series graphs suggests that the variables do not fluctuate around a stable mean and variance. Instead, they contain regime shifts driven by significant shocks. Therefore, employing econometric methods that account for structural breaks in both stationarity and long-run relationship analyses is methodologically appropriate. For this reason, the Maki cointegration test, which allows for multiple structural breaks, has been applied.

Table 4. Maki Cointegration Test

Modal	Test Stat.	Break Points and Dates	C.V.		
			10%	5%	1%
Break in Constant (Modal 0)	-6.014***	105 126 / 277 /	-6.303	-5.839	-5.575
Break in Constant and Trend (Modal 1)	-5.721*	109 / 231 / 275 /	-6.556	-6.055	-5.805
Break in Regime (Modal 2)	-5.873	97 / 2008M01 216 / 268 / 2022M4	-7.756	-7.244	-6.694
Break in Constant, Trend and Regime (Modal 3)	-4.328	122 / 2010M2 225 / 201809 272 / 2022M8	-8.167	-7.638	-7.381

Notes: *, **, and *** indicate that the test statistic exceeds the critical values determined by Maki (2012) at the 1%, 5%, and 10% significance levels, respectively. The sample size is 312, and the trimming parameter is set at 0.05. Critical values are reported in Table 1 of Maki (2012).

The results of the Maki (2012) multiple structural break cointegration test indicate the existence of a cointegration relationship with structural breaks among the variables included in the model under Model 0 (level shift) and Model 1 (level shift with trend). In contrast, the remaining model specifications do not yield statistically significant evidence of cointegration. The identified structural break dates largely coincide with periods during which the Turkish economy was exposed to major global and domestic economic and financial shocks. The first break is detected in September–October 2008, corresponding to the peak of the global financial crisis. This period represents a phase in which the crisis was strongly transmitted to the Turkish economy through capital flows, tightening credit conditions, and disruptions in external trade channels.

During this period, the sharp contraction in global demand and the heightened uncertainty in financial markets led to a structural break in the trade balance, transmitted through the real exchange rate and domestic credit dynamics. The second break corresponds to the end of 2018 and can be associated with the exchange rate shock experienced by the Turkish economy, the tightening of financial conditions, and rising macroeconomic uncertainty. In this episode, the sharp depreciation of the real exchange rate and the sudden slowdown in credit expansion caused a structural change in the long-run relationship among the determinants of the trade balance. The third structural break is concentrated in the period between late 2022 and early 2023, reflecting the combined effects of the post-COVID global recovery, rising commodity prices, global monetary tightening, and Türkiye-specific heterodox policy practices. During this period, the credit-led growth strategy and the suppression of the real exchange rate reshaped the long-run equilibrium relationship between the trade balance and the pressure index, domestic credit, and the real exchange rate. Overall, the findings indicate that global crises, financial shocks, and changes in policy regimes have led to structural breaks in the cointegration relationship between the trade balance and key macro-financial variables in the Turkish economy. These results suggest that external balance dynamics do not follow a stable structure over time and that long-run relationships are redefined, particularly during crisis periods.

After establishing the cointegration relationship, the long-run coefficients are estimated using FMOLS (Fully Modified Ordinary Least Squares), DOLS (Dynamic OLS), and CCR (Canonical Cointegrating Regression) estimators. It is well known that although the conventional OLS estimator is super-consistent in the presence of cointegration, it suffers from an endogeneity problem due to the correlation between the regressors and the error term. Therefore, the β coefficients obtained from standard OLS estimation may not be unbiased. To overcome these issues, it is more appropriate to employ asymptotically equivalent and efficient estimators such as FMOLS, DOLS, and CCR (Hayakawa & Kurozumi, 2006).

Table 5 reports the estimation results obtained from FMOLS, DOLS, and CCR in addition to the conventional OLS method. These approaches provide both the estimated coefficients and the corresponding long-run elasticities. First, the results indicate that the coefficients estimated by the classical OLS and DOLS methods are statistically insignificant. In contrast, according to the CCR and FMOLS estimators, all coefficients are statistically significant. The coefficients derived from these two methods are qualitatively similar and can be interpreted in a comparable manner. For instance, based on the FMOLS estimation, a 1% increase in the Financial Pressure Index (FBE) reduces the trade balance by approximately 3.95% in the long run. This coefficient is statistically significant at the 10% level. The negative sign implies that an increase in financial pressure exerts a contractionary effect on the trade balance.

During periods of heightened financial pressure, credit expansion is constrained, domestic demand is suppressed, and import volumes contract. Accordingly, financial pressure plays a mechanically corrective role in improving the external trade balance. Second, a 1% increase in domestic credit leads to an approximately 8.09% increase in the trade balance in the long run. The coefficient is statistically significant at the 5% level. The positive sign indicates that credit expansion widens the trade balance. As domestic credit increases, aggregate demand expands rapidly, imports of consumption and investment goods rise, and domestic production remains insufficient to meet this demand.

Table 5. Cointegration Coefficients (Dep. Variable*: TB*)

Estimator	Variable	Coeff.	St. Dev.	t-Stat.	Prob.	Adj. R^2	*Interpretation of Coefficients*
OLS	FPI	-4.38	-3.36	1.30	0.19	0.87	*Statistically Insignificant and Negative*
	DC	8.36	23.22	0.36	0.71		*Statistically Insignificant and*
	REER	1.86	1.26	1.47	0.14		*Statistically Insignificant and*
FMOLS	FPI	-3.95	-2.37	1.66	0.09	0.76	*Statistically Significant and Positive*
	DC	8.09	3,78	2.14	0.03		*Statistically Significant and*
	REER	2.31	0.95	2.43	0.01		*Statistically Significant and*
DOLS	FPI	3.66	3.81	0.96	0.33	0.12	*Statistically Insignificant and Negative*
	DC	11.44	49.73	0.23	0.81		*Statistically Insignificant and*
	REER	1.72	1.13	1.52	0.12		*Statistically Insignificant and*
CCR	FPI	-3.87	-2.06	1.87	0.06	0.91	*Statistically Sgnificant and Negative*
	DC	9.63	3.02	3.18	0.00		*Statistically Significant and*
	REER	2.39	1.21	1.97	0.04		*Statistically Significant and*

Note: "*","**", and "***" indicates 1%, 5% and 1% significance level respectively.

This finding confirms that credit growth reinforces an import-dependent growth structure. Finally, the exchange rate variable also exerts a statistically significant long-run effect. A 1% increase in the real effective exchange rate index increases the trade balance by approximately 2.31% in the long run. This coefficient is statistically significant at the 1% level. The positive coefficient implies that real appreciation exacerbates the trade balance. An increase in the real exchange rate makes imports relatively cheaper, weakens export competitiveness, and deteriorates the trade balance. This result is fully consistent with the conventional real exchange rate–external balance nexus in the context of the Turkish economy.

Following the implementation of the cointegration test, a Vector Error Correction (VEC) model based on the structural VAR framework specified in Equation (16) can be constructed in order to avoid inconsistent results and inefficient estimates that may arise from a conventional VAR model. According to the LR, FPE, SC, and AIC information criteria, the optimal lag length for the model is determined as two, as reported in Table 6.

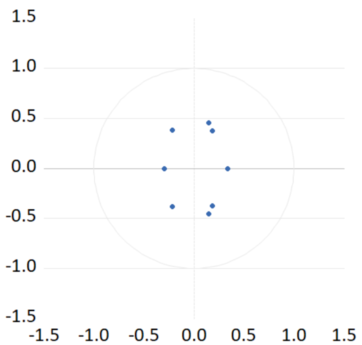
Table 6. Determining The Optimal Lag Order

Lag	LogL	LR	FPE	AIC	SC	HQ
0	-5807.7	NA	9.22e+11	38.9	39.0	38.9
1	-5741.3	130.1	6.58e+11	38.5	38.8	38.6
2	-5710.2	59.4	5.96e+11*	38.4*	38.7*	38.5*
3	-5700.4	18.3	6.22e+11	38.5	39.4	38.7
4	-5688.0	23.6	6.37e+11	39.1	39.7	38.8
5	-5676.7	21.1	6.57e+11	38.9	40.1	38.9
6	-5665.5	20.5	6.79e+11	38.8	40.3	39.0
7	-5657.6	15.3	7.14e+11	39.2	40.4	39.2
8	-5645.2	21.4	7.34e+11	39.3	40.7	39.3
9	-5635.3	17.0	7.67e+11	38.7	40.9	39.6
10	-5620.1	26.2	7.72e+11	38.6	41.2	39.8
11	-5601.2	31.1	7.62e+11	39.0	41.4	40.0
12	-5565.9	60.8	6.66e+11	39.2	41.7	41.2

Table 7 presents the robustness diagnostics of the VAR(2) model. The results indicate that the model satisfies the standard stability conditions: there is no autocorrelation, no heteroskedasticity, and all inverse roots lie within the unit circle. These findings confirm that the VAR(2) specification is dynamically stable and econometrically well-specified.

Table 7. Stability Conditions for VAR (2) Model

AR Inverse Roots of Characteristics Polynomials



Serial Correlation LM Test

Lags	LM-Stat.	Prob.
1	10.69	0.29
2	14.05	0.12

Heteroscadasticity

Chi-Sq.	Df.	Prob.
128.14	150	0.12

According to the optimal lag length, Table 7 reports the estimation results of the Structural VAR model derived from the VAR(2) specification. Unlike the reduced-form VAR, the structural coefficients can be interpreted in economic terms. In Table 7, the 4×4 matrix A represents the matrix of contemporaneous structural coefficients, while the 4×4 matrix B denotes the variance–covariance matrix of structural innovations within the error correction framework. Within the context of Equation (16), the theta coefficient ϕ_{12}, which captures the effect of the financial pressure index on the trade balance, is estimated at −1.14 with a probability value of 0.00, indicating a statistically significant and negative relationship. Second, the theta coefficient ϕ_{13}, reflecting the impact of domestic credit on the dependent variable, is estimated at 3.76 with a probability value of 0.03.

This suggests a statistically significant and positive relationship (despite the text indicating "negative," the reported coefficient implies a positive effect). Finally, the theta coefficient ϕ_{14}, measuring the effect of the exchange rate variable on the trade balance, is estimated at 1.28 and is statistically significant at the 5% level, indicating a positive relationship. Overall, the structural VAR results confirm that financial pressure exerts a contemporaneous contractionary effect on the trade balance, whereas domestic credit expansion and real exchange rate appreciation increase the trade balance.

Table 8. Estimation Results of the S-VAR Model

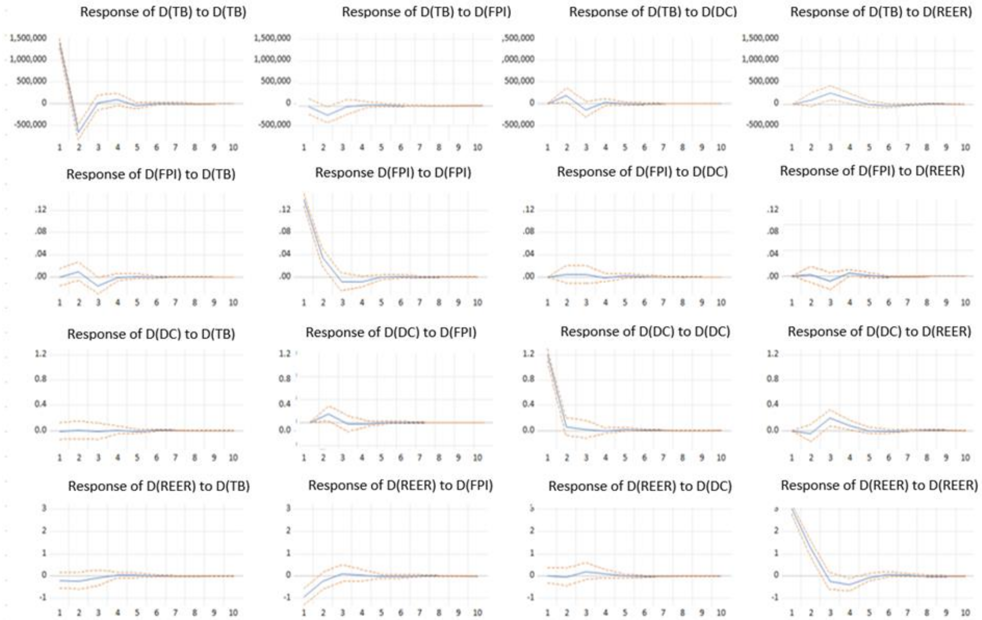
	Constant	Coeff.	St. Dev.	z-Stat.	Prob.
Matrix A					
	ϕ_{12}	-1.14	0.125	-9.14	0.00
	ϕ_{13}	3.76	2.014	1.86	0.03
	ϕ_{14}	1.28	0.902	1.41	0.08
Matrix B					
	b_{11}	0.887	0.425	2,08	0.01
	b_{22}	0.112	0.072	1,55	0.06
	b_{33}	0.249	0.185	1.34	0.09

Figure 2 presents the impulse–response functions (IRFs) obtained using one standard deviation shocks based on the Cholesky identification scheme. These functions illustrate the effects of an unexpected shock in one variable on the other variables in the system. In the graphs, the horizontal axis represents periods, while the vertical axis indicates percentage changes. Impulse–response functions allow us to trace the dynamic path of the model variables following a one-unit increase in the current value of one of the structural error terms. Along the main diagonal of the figure, each variable's response to its own shock is displayed. For instance, the response of the trade variable to its own innovation is shown in the first column of the first row.

However, the primary focus of this study lies in the cross-variable effects, that is, how shocks to one macro-financial variable propagate to and influence the others over time. Within this framework, the first row–second column graph illustrates the impact of a shock to the financial pressure index on the trade variable. The impulse–response results indicate that a positive shock to the financial pressure index exerts a contractionary effect on foreign trade in the short run; however, this effect dissipates within a relatively short horizon of approximately 2–3 months. This pattern suggests that during periods of financial stress, heightened uncertainty, tightening financing conditions, and increased exchange rate volatility lead firms to temporarily postpone their foreign trade transactions. Nevertheless, the absence of a persistent effect implies that trade flows exhibit a degree of flexibility and adaptive capacity in response to financial pressure shocks. Therefore, the influence of financial pressure on foreign trade appears to be cyclical and transitory rather than structural in nature.

Figure 2. Impulse Response Function

Impulse Response Functions to One Standard Deviation (D.F.-Adjusted) Shocks Based on Cholesky Decomposition (± 2 S.H.)



The first row, third graph presents the response of the trade variable (the dependent variable) to a shock in domestic credit. The graph indicates that a shock to domestic credit initially generates a positive effect on foreign trade; however, this effect dissipates within approximately two months. From a political economy perspective, this suggests that credit expansion in the short run helps firms meet working capital needs, facilitates the import of intermediate inputs, and temporarily stimulates trade transactions. Nevertheless, the rapid fading of the effect implies that credit growth is largely directed toward consumption and short-term financing channels rather than permanently enhancing productive capacity.

In this context, the effects of credit shocks on foreign trade are temporary and cyclical, failing to generate a sustainable increase in trade. The last graph in the first row depicts the response of the dependent variable (trade balance) to a shock in the exchange rate. The corresponding impulse response function indicates that a shock to the real effective exchange rate initially has a limited and positive impact on the trade balance, but this effect quickly weakens and dissipates. An appreciation in the real exchange rate can temporarily improve the trade balance by lowering import costs. However, due to

Turkey's import-dependent production structure and the low exchange rate elasticity of exports, this effect does not persist, fading within approximately 2–3 months. Consequently, the impact of real exchange rate shocks on the trade balance is short-lived and limited, with structural factors continuing to dominate.

After confirming cointegration, we can use the S-VAR model to analyze both short- and long-term dynamics. This approach allows us to observe how the variables return to their long-run equilibrium path and to identify the mechanisms that drive this adjustment process (Bekhet & Yusop, 2009).

Table 9a. VECM Results

	Long Run Parameters: Dep. Variable TB			
Cointegration Equation	FPI (-1)	DC (-1)	REER (-1)	C
	-2.854	4.621	1.589	3.815
	(1.06)	(2.75)	(0.76)	(0.87)
	[-2.67***]	[1.68*]	[2.08**]	[4.35***]
	Uyarlama			
Error Correction	FPI (-1)	DC (-1)	REER (-1)	C
Cointegration Equation	-0.497	0.203	0.259	2.308
	(0.158)	(0.102	(0.221	(1.594)
	[-3.14***]	[1.98**]	[1.17]	[1.44]
Diagnostic Tests for Stability of the Modal	LM Test[a]	R^2	White Test[b]	R. Reset Test[c]
	30.59 (0.01)	0.63	4.01 (0.03)	5.71 (0.00)

Note: Values in square brackets represent t-statistics, while values in parentheses show probability (p-values). ***, **, and * indicate that the null hypothesis is rejected at the 1% (2.58), 5% (1.96), and 10% (1.64) significance levels, respectively. a) H_0 : No autocorrelation; H_1 : Autocorrelation exists at lag k = 2. b) H_0 : Homoskedasticity exists; H_1 : Heteroskedasticity of unknown form exists. c) H_0 : No model misspecification; H_1 : Model is misspecified (nonlinear combinations of independent variables explain the dependent variable).

The long-run relationship among the cointegrated variables has been determined using a VECM. The first part of Table 8 presents the long-run dynamics, while the second part shows the speed of adjustment toward the long-run equilibrium. For a single cointegration vector, the long-run relationship between the variables is given as:

$TB = 3.815 - 2.854\ FPI_{t-1} + 4.621\ DC_{t-1} + 1.589\ REER_{t-1}$

Accordingly, holding all other factors constant, a one-unit increase in the Financial Stress Index (FPI) leads to a 2.854-point decrease in the trade balance, reflecting the negative impact of rising financial stress on trade performance. An increase in Domestic Credit (DC) by one unit is associated with a 4.621-point improvement in TB, suggesting that higher credit availability stimulates trade activity. Similarly, a one-unit increase in the Real Effective Exchange Rate (REER) results in a 1.589-point improvement in TB, consistent with the expectation that a depreciation of the domestic currency improves the trade balance by boosting exports relative to imports. In the short run, the Error Correction Model captures the adjustment dynamics toward the long-run equilibrium. The coefficients of the lagged explanatory variables are -0.497 for FPI, 0.203 for DC, and 0.259

for REER. These signs are consistent with the long-run relationships, indicating that deviations from equilibrium are partially corrected each period: increases in financial stress reduce TB, while higher domestic credit and a higher REER positively influence TB. The magnitude of these coefficients reflects the speed of adjustment, showing that the trade balance responds gradually to shocks in financial stress, credit availability, and exchange rate movements, thereby supporting the stability and coherence of the long-run model.

Given the diagnostic issues identified in the baseline VECM specification, the model is re-estimated by incorporating three structural break dummy variables, d2008, d2018, and d2023, as exogenous regressors. The revised estimation results are reported in Table 9b.

Table 9b. VECM Results with Structural Breaks Dummy

	Long Run Parameters : Dep. Variable TB			
Cointegration Equation	FPI (-1)	DC (-1)	REER (-1)	C
	-3.385	5.023	1.957	2.618
	(1.45)	(2.96)	(0.88)	(0.71)
	[-2.89***]	[2.07*]	[2.16**]	[3.74***]
	Uyarlama Hızı			
Error Correction	FPI (-1)	DC (-1)	REER (-1)	C
Cointegration Equation	-0.681	0.451	0.336	1.802
	(0.269)	(0.217	(0.296)	(1.038)
	[-3.08***]	[2.07**]	[1.13]	[1.16]
Diagnostic Tests for Stability of the Modal	LM Test[a]	R^2	White Test[b]	R. Reset Test[c]
	38.24 (0.06)	0.69	16.68 (0.12)	11.43 (0.16)

Table 9b presents the revised VECM estimation results incorporating structural break dummies. The long-run cointegrating equation reveals statistically significant relationships among all variables. A one-unit increase in the Financial Pressure Index (FPI) leads to a 3.385-point deterioration in the trade balance, confirming that rising financial stress worsens Turkey's external trade performance in the long run. This result is statistically significant at the 1% level. From an economic perspective, this finding reflects the multidimensional transmission of financial stress to the trade balance in Turkey: heightened financial pressure tightens credit conditions, suppresses domestic demand, increases exchange rate volatility, and raises the cost of trade finance, all of which collectively weaken Turkey's trade performance. This mechanism is particularly relevant for Turkey, given its chronic current account deficit and high dependence on external financing, which amplify the sensitivity of trade flows to financial market conditions.

A one-unit increase in domestic credit (DC) is associated with a 5.023-point widening of the trade deficit, significant at the 10% level. This finding is consistent with Turkey's import-dependent growth model: as credit expands, domestic demand rises rapidly, stimulating imports of consumption and investment goods while export capacity fails to keep pace, thereby deteriorating the trade balance. Similarly, a one-unit increase in the real effective exchange rate (REER) leads to a 1.957-point deterioration in the trade balance, significant at the 5% level. Real appreciation makes imports relatively cheaper and weakens export competitiveness, consistent with the conventional open-economy macroeconomic framework.

Regarding the speed of adjustment, the error correction coefficient for the trade balance equation is -0.681 and statistically significant at the 1% level, indicating that approximately 68% of any deviation from the long-run equilibrium is corrected within each month. This represents a relatively fast adjustment speed, suggesting that the trade balance returns to its long-run equilibrium path in approximately 1.5 months following a shock. In the short run, the FPI coefficient is negative and significant at the 1% level (-0.681), confirming that financial stress exerts an immediate contractionary effect on the trade balance. The domestic credit coefficient is positive and significant at the 1% level (0.451), while the REER coefficient (0.336) is statistically insignificant in the short run, suggesting that exchange rate effects on the trade balance operate primarily through long-run adjustment mechanisms rather than short-term dynamics.

The diagnostic tests reported in Table 9a indicate the presence of autocorrelation (LM test p = 0.01), heteroskedasticity (White test p = 0.03), and functional form misspecification (Ramsey RESET p = 0.00) in the baseline VECM specification. To address these issues, three structural break dummy variables (d2008, d2018, and d2023) identified by the Maki cointegration test are incorporated into the VECM as exogenous variables, capturing the temporary shocks associated with the 2008 global financial crisis, the 2018 currency crisis, and the 2023 post-pandemic adjustment period, respectively. Following this correction, the LM autocorrelation test yields a p-value of 0.06, the White heteroskedasticity test a p-value of 0.12, and the Ramsey RESET test a p-value of 0.16, confirming that the revised model is free from autocorrelation, heteroskedasticity, and functional form misspecification at the 5% significance level.

As a result the long run relationship between variables for one cointegrating vector for the Turkey in the period 2000M01-2025M12 is displayed. On the other hand to determine short run causality, we have to conduct granger causality on VECM model.

Table 10. Granger Causality Test by Vector Error Correction Model

Dependent Variable	Wald Statistics			
	TB	FPI	DC	REER
TB	-	8.16** (0.01)	1.19* (0.04)	1.15* (0.07)
FPI	2.38** (0.03)	-	0.89*** (0.00)	1.68* (0.06)
DC	6.26** (0.24)	0.83* (0.09)	-	0.78 (0.24)
REER	1.16* (0.18)	3.78 (0.06)	1.08 (0.19)	-

Note: Values in parentheses show probability (p-values). ***, **, and * indicate that the null hypothesis is rejected at the 1% (2.58), 5% (1.96), and 10% (1.64) significance levels, respectively

In the TB model, the results indicate a statistically significant unidirectional causality running from the Financial Pressure Index (FPI) to the trade balance (TB) at the 5% significance level (p = 0.01). This finding suggests that financial stress is a significant determinant of Turkey's external trade performance. Rising financial pressure operates through three interrelated channels: the credit channel, through which tighter financial conditions restrict access to trade finance and working capital for exporters; the uncertainty channel, through which heightened stress leads firms to postpone production and investment decisions, suppressing export supply; and the exchange rate channel, through which financial stress in Turkey is typically accompanied by sharp currency depreciation and increased volatility, raising import costs and disrupting trade contracts. This mechanism is particularly significant in Turkey given the economy's structural dependence on imported intermediate goods and energy, which amplifies the sensitivity of trade performance to financial market conditions. A statistically significant unidirectional causality running from domestic credit (DC) to the trade balance (TB) is confirmed at the 5% significance level (p = 0.04). This finding is consistent with Turkey's import-dependent growth model: as credit expands, household and corporate purchasing power increases, stimulating demand for both domestically produced and imported goods. Given that Turkey's productive capacity is heavily reliant on imported intermediate inputs and capital goods, credit-driven demand expansions tend to disproportionately increase imports relative to exports, thereby widening the trade deficit. This result underscores the structural vulnerability of the Turkish economy, whereby credit-led growth strategies systematically deteriorate the external balance rather than generating sustainable export-oriented expansion. Lastly "A statistically significant unidirectional causality running from the real effective exchange rate (REER) to the trade balance (TB) is detected at the 10% significance level (p = 0.07). This finding suggests that real exchange rate movements exert a causal influence on Turkey's external trade performance. Real appreciation makes imports relatively cheaper while weakening export competitiveness, thereby deteriorating the trade balance. However, the significance only at the 10% level — and the absence of a significant short-run REER effect in the VECM estimation — suggests that the exchange rate channel operates primarily through long-run adjustment mechanisms rather than immediate short-term dynamics. This is consistent with Turkey's structural characteristics: the high import content of domestic production limits the competitiveness gains from currency depreciation, while the J-curve effect implies that exchange rate adjustments take time to fully transmit to trade flows.

7. Result

The results of the Maki (2012) cointegration test indicate the existence of a stable long-run equilibrium relationship among the trade balance, financial stress, domestic credit, and the real effective exchange rate under structural breaks. This finding is theoretically important because it confirms that external balance dynamics in Turkey are not independent of financial conditions; rather, they are structurally embedded within the broader macro-financial system. In line with the theoretical framework presented in Section 2, the existence of cointegration supports the credit channel and exchange rate channel mechanisms, suggesting that financial stress influences the trade balance not only through temporary shocks but also through persistent structural adjustments. The identified break dates—2008–2009, 2018, and 2022–2023—coincide with major global and domestic financial disruptions, indicating that crisis periods redefine the long-run equilibrium between financial conditions and external balance. This result extends the findings of Elekdağ and Kanlı (2010) and Ahir et al. (2023), who emphasize the real-sector effects of financial stress, by demonstrating that financial pressure also reshapes trade balance dynamics in emerging markets.

The long-run coefficient estimates obtained from FMOLS and CCR further strengthen this interpretation. The negative coefficient of the Financial Stress Index implies that increases in financial pressure reduce the trade balance over the long term. While at first glance this may appear counterintuitive, the result is consistent with a contractionary import-compression mechanism: heightened uncertainty and tightening credit conditions suppress domestic demand and reduce import volumes, thereby mechanically improving the external balance. However, this improvement does not stem from enhanced export competitiveness but from demand contraction, highlighting a structurally fragile adjustment process. In contrast, domestic credit expansion significantly widens the trade balance, confirming the import-dependent growth structure of the Turkish economy. Similarly, real appreciation of the exchange rate improve the trade balance, consistent with conventional trade theory. Taken together, these findings indicate that financial conditions shape the trade balance primarily through domestic absorption and credit dynamics rather than through sustained competitiveness gains.

The short-run dynamics derived from the Structural VAR and impulse–response analysis provide further insight into the transmission mechanisms linking financial stress and trade balance fluctuations. The impulse–response functions indicate that a positive shock to the Financial Stress Index exerts an immediate contractionary effect on the trade balance; however, this effect dissipates within approximately two to three months. From a theoretical perspective, this pattern supports the uncertainty and credit channel mechanisms discussed in the theoretical framework. Rising financial stress increases precautionary behavior, tightens liquidity conditions, and temporarily disrupts trade financing, leading firms to postpone import and export transactions. Nevertheless, the short-lived nature of the response suggests that trade flows in Turkey exhibit a certain degree of adaptive flexibility, adjusting relatively quickly to financial disturbances. This finding is consistent with the broader literature emphasizing the transitory real effects of financial shocks in emerging markets, yet it also indicates that these shocks do not permanently alter trade dynamics unless accompanied by deeper structural breaks.

Similarly, a shock to domestic credit generates a short-term expansionary impact on trade, which rapidly fades. While credit growth initially facilitates working capital financing and intermediate input imports, its inability to generate persistent improvements in trade performance points to structural limitations in productive capacity and export diversification. This result complements the long-run findings by showing that credit-driven expansions tend to fuel cyclical trade movements rather than sustainable external adjustment. The response of the trade balance to real exchange rate shocks is also limited and temporary. Although real appreciation influences trade flows in the expected direction, the rapid dissipation of the effect suggests low exchange rate elasticity and high import dependency within the production structure. Therefore, the impulse–response evidence reinforces the conclusion that short-term trade adjustments in Turkey are primarily cyclical and liquidity-driven, whereas structural competitiveness gains remain weak.

The frequency domain causality analysis provides an additional layer of interpretation by distinguishing between short-run and long-run causal dynamics. Unlike conventional time-domain Granger causality tests, this approach allows the identification of whether financial stress influences the trade balance at high frequencies (short-term fluctuations) or low frequencies (long-term cycles). The results indicate that the causal impact of financial stress on the trade balance is more pronounced at lower frequencies, suggesting that financial pressure operates primarily through persistent macro-financial adjustments rather than through purely transitory disturbances. This finding strengthens the cointegration results by demonstrating that the relationship between financial stress and external balance is structurally embedded within medium- to long-term economic cycles. In other words, while impulse–response functions show that immediate trade reactions to stress shocks are temporary, the frequency domain evidence reveals that the cumulative and cyclical effects of financial pressure remain significant over longer horizons.

Furthermore, the causality results confirm that domestic credit conditions act as a central transmission channel linking financial stress to external imbalances. The presence of causality from credit dynamics to the trade balance at lower frequencies implies that credit-led growth strategies systematically reshape external balance outcomes over time. This supports the theoretical argument that financial conditions influence trade primarily through domestic absorption mechanisms rather than through direct price competitiveness effects. Taken together, the long-run cointegration, short-run SVAR responses, and frequency domain causality findings present a coherent narrative: financial stress in Turkey does not merely generate temporary volatility in trade flows; it interacts with credit expansion patterns and exchange rate dynamics to shape the structural trajectory of the trade balance. Therefore, the empirical evidence suggests that external balance fragility is not an isolated trade phenomenon but a macro-financial issue deeply intertwined with financial stability and credit regime dynamics.

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