



The regime switching evidence of financial-economic-political risk in Turkey

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Abstract

In recent time, Turkey could be said to have experienced different levels of Economic Risk, Financial Risk, and Political Risk from low- to high-level. This study investigates the linkage between country risks, namely Financial Risk, Economic Risk, and Political Risk (FEP risk) in Turkey for the period 1984Q1 to 2019Q1 by using threshold cointegration, Markow-switching regression (given the nonlinearity and structural breaks observed in the time series variables), and frequency domain causality approaches. The empirical findings of this study reveal that (i) nonlinear cointegration between Economic Risk, Financial Risk, and Political Risk in Turkey is statistically significant given the evidence of threshold cointegration test, which determines the structural breaks endogenously; (ii) there is positive linkage among the component of country risk at different volatility periods; (iii) there is a significant Granger causal linkage between Economic Risk, Financial Risk and Political Risk at the different frequency levels. The study is likely to open debate about the literature since the study concludes with a discussion on short-run and long-run implications for economic, political, and financial stabilises, thus offering policy suggestions for the policymakers in Turkey.

Keywords Economic risk · Financial risk · Political risk · Threshold cointegration · Frequency domain causality · Turkey

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1 Introduction

Since the global crisis in 1929, researchers have given considerable attention to the impact and consequence of the political, financial, and economic crises. In the relevant economic literature, although numerous empirical and theoretical studies have investigated (i) the nexus between economic growth and finance development and (ii) the effect of political instability on economic growth. However, two-way relationship between country risk components, namely Economic Risk, Financial Risk, and Political Risk is yet to receive considerable attention and especially for the case of Turkey. Although Turkey has continued to experience economic setback in recent times arising from high inflation and currency turmoil, the country experienced significant transformation since the late 1980s by liberalizing their markets, accelerating reforms, the opening of the Turkish economy to liberal trade and adopting modern banking systems. Therefore, the main innovation of the present study is to construct time series-based models to explore the relationship between Economic Risk, Financial Risk, and Political Risk which has not widely been investigated for the case of Turkey especially by exploring more recently developed econometrical tools, namely threshold cointegration, Markow-switching regression and frequency domain causality tests which takes into account structural breaks. Thus, the study is billed to the fill gap in the literature at a timely manner given Turkey's currency and exchange rate turmoil alongside affirming the appropriateness of the time series-based models. Therefore, by opening a a new debate in the literature, the findings of this investigation expectedly brings out a remarkable suggestion for policymakers in Turkey.

2 Study relevance and literature

The vulnerabilities in the Turkish economy and political environments have encourages researchers to ponder and investigate potential causals and its associated effects. Chronologically, the Turkish economy is dramatically affected by the 1988 stock market and currency crisis, the 1994 economic crisis, the 1998 textile crisis, the 2000 banking crisis, the 2001 economic crisis, the 2007–08 global economic crisis, the 2018 exchange rate crisis. As mentioned previously, in the last decade, Turkey has experienced economic setbacks characterized by rising inflation, currency depreciation among other monetary policy issues, not mentioning the fact that recent development has also shown the country's vulnerability to political tension. It has witnessed a mass uprising characterized by several protests at Istanbul's Taksim Gezi Park in 2013 and for other reasons such as the local election in March 2014, a presidential election in August 2014, two general elections in June and November 2015, a failed coup attempt in July 2016, the declaration of a state of emergency in that same month, and finally, the April 16, 2017 referendum" (Akçay 2018).

2.1 Related literature

It is widely accepted that vulnerability in the political system is likely to destabilize macroeconomic dynamics. This hypothesis was initially put forward by the study of Olson (1963), which underlined that political stability is a crucial condition for the development and encouragement of entrepreneurs and forecasting a nation's long-term economic performance. Therefore, the scholars- Brunetti (1997), Chen and Feng (1996),

Jong-a-Pin (2009), Alesina and Perotti (1996), Darby et al. (2004), Aisen and Veiga (2013)- have constructed ample literature documenting the adverse impact of political uncertainty on macroeconomic dynamics, namely GDP growth, private investment, taxation, public expenditures, and investment, debt, and inflation. Also, the financial system and market are adversely affected by rising political tension in a country. Cutler, Poterba, and Summers (1989) and Hibbs (1986) argue that government policy changes lead to changes in the stock market returns. Pantzalis et al. (2000), Li and Born (2006), and Bialkowski, Gottschalk, and Wisniewski (2008) explored the behaviors of asset prices in the stock markets for the political elections. They concluded that the price of stock market assets is significantly affected by political elections. However, in the literature, the reverse relation has not been comprehensively investigated, including the economic/finance-led politic hypothesis.

Regarding the channels of transmission, Campos and Nugent (2002) argue that slowing down economic development is likely to change governments' structure due to changing the power balance in the government. Thus, political instability in a country might be raised due to the collapse of the government. The finding of Telatar (2003) also supports the argument of Campos and Nugent (2002). By focusing on 122 countries, MilijkoVIC and Rimal (2008) put it forward that social, economic conditions such as income level lead to changes in the political system. Therefore, political vulnerability is likely to be triggered by a vulnerability in economic growth.

Although numerous empirical and theoretical studies have been explored the nexus between economic growth and financial development in the relevant economic literature, there is no consensus about the direction of the widespread phenomenon of the relationship between financial development and economic growth since the finance-led growth hypothesis put it forward by Schumpeter (1912). In examining the linkage between economy and finance, scholars have mainly tested three different hypotheses, namely "finance lead growth," "growth lead finance", and "feedback," to be able to understand their causal natures. Financial development is the fundamental tool for economic development as financial institutions could accelerate economic growth via improving risk management, making financial transactions, exchanging goods and services, and saving mobility easier (Levine et al. 2000). This hypothesis has been supported by the well-documented studies of King and Levine (1993), Levine et al. (2000), McKinnon (1973), Rajan and Zingales (1998), Galindo et al. (2007), Marashdeh and Al-Malkawi (2014), Cournède and Denk (2015), Cojocarui et al. (2016)-. According to growth led finance hypothesis, a well-functioning financial system is the consequence of economic growth. Robinson (Robinson, 1952) was the pioneer economist who states that finance is not an important factor in the growth process, but economic growth may boost the financial industry. The studies of Shaw and Gurley (1967), Stiglitz (1994), Zang and Kim (2007), Odhiambo (2008), Jenkins and Katircioglu (2010) also underlined the importance of economic growth on financial development. Lastly, the third hypothesis argues that there is a bidirectional causal relationship between financial development and economic growth (Blackburn and Hung 1998; Liang and Jian-Zhou 2006). A few studies also supported the neutrality hypothesis and showed no causal relationship between these two financial and economic growth variables (Lucas 1988; Stern 1989; Nyasha and Odhiambo 2015). Moreso, recent studies have also linked economic policy and its aspects to other socio- and economic indicators such as energy and environment (Anser et al., 2021; Adams et al., 2020; Akadiri et al., 2020).

Exploring the relationship between Economic Risk, Financial Risk, and Political Risk in Turkey is the paper's primary objective. It is crucial for governors, investors

Table 1 Data and descriptive statistics

Data	Economic risk index	Financial risk index	Political risk index
Source	PRS	PRS	PRS
Code	ERI	FRI	PRI
Mean	27.576	30.593	56.643
Median	27.833	31.666	56.500
Maximum	36.000	38.500	70.000
Minimum	13.280	22.333	41.666
Std. Dev	6.799	4.193	6.251
Skewness	-0.371	-0.404	0.045
Kurtosis	1.803	2.192	2.853
Jarque-Bera	11.646	7.674	0.174
Probability	0.002	0.021	0.916

and also academics. Still, to the best of our knowledge, specifically, there is no study of Turkey in the literature that have examined the aforementioned relationship especially by using threshold cointegration, Markow-Switching regression, and frequency domain causality tests simultaneously. As its main innovation, the present study constructs a time series-based model to fill this literature gap. The present study is likely to open a new debate in the literature, and the findings bring out the noteworthy implication for governors in Turkey.

3 Data description and methodology

This study employs data sets of Political Risk Index, Economics Risk Index, and Financial Risk Index as variables of concern for Turkey's case. The datasets that span 1984Q1 to 2019Q1 were retrieved from the Political Risk Services (PRS) Group. As indicated below, the categorical description of Political Risk Index, Economics Risk Index, and Financial Risk Index are harmonized explicitly as:

- The Political Risk Index: "It is a country type of risk associated with the vulnerabilities and changes in the political dynamics or structure of a state. Political risk could be traced to the unusual temperate and uncertainty in government stability, bureaucratic quality, internal and external conflicts, religious and ethnic problems, and others. The PRS Group assesses the Political Risk Index on a scale of 0 (maximum risk) to 100 (minimum risk)."
- The Economic Risk Index: "it measures the economic weaknesses and strengths, with an assigned value between 0 and 50. While the value 0 corresponds to the highest economic risk, the lowest economic risk is assigned 50. The assessment of the PRS Group is based on the GDP per capita, GDP growth, inflation, and the current account as a percentage of GDP, along with the budget balance as a percentage of GDP variables."
- The Financial Risk Index: "This measures the country's ability to pay its debts in general. The variable risk range is between 0 (maximum risk) and 50 (minimum

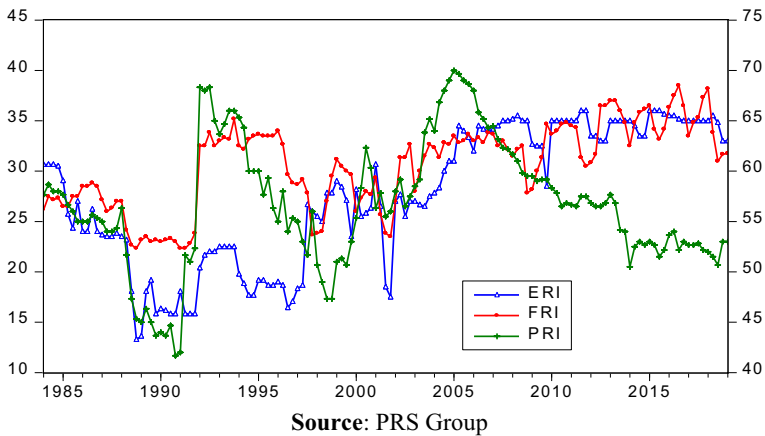


Fig. 1 Financial Risk, Economic Risk and Political Risk in Turkey Source: PRS Group

Table 2 Zivot Andrew unit root test

	ERI	FRI	PRI
<i>Series in levels</i>			
C	-4.375 (2002Q1)**	-3.162 (1992Q1)*	-4.078 (1991Q2)**
C&T	-4.0981 (1997Q3)**	-3.530 (1992Q1)	-4.399 (1991Q2)**
<i>Series in First Differences</i>			
C	-9.666** (1990Q1)	-5.800** (2002Q1)**	-5.932** (1991Q1)
C&T	-7.199** (1999Q2)	-6.074** (1993Q2)**	-6.503** (1992Q4)**

Note: C and C&T denote constant and constant and trend in the ZA unit root test, respectively. ** and * indicate statistical significance at 0.05 and 0.10 levels, respectively. The numbers in parenthesis () represent breakpoints

risk). The exchange rate, liquidity variables, foreign debt as a percentage of GDP, current account as a percentage of goods and services are among many of the variables utilized in assessing the Financial Risk Index."

While Table 1 presents the descriptive statistics of the variables of interest and their respective codes, Fig. 1 visually indicates the potential breaks and patterns of the time series variables.

Table 3 BDS dependency test

	Dimension	BDS statistic	z-Statistic	Prob
ERI	2	0.161	40.062	0.000**
	3	0.269	42.005	0.000**
	4	0.336	43.961	0.000**
	5	0.376	47.139	0.000**
	6	0.398	51.515	0.000**
FRI	2	0.145	33.242	0.000**
	3	0.236	34.425	0.000**
	4	0.288	35.597	0.000**
	5	0.313	37.514	0.000**
	6	0.323	40.618	0.000**
PRI	2	0.155	25.023	0.000**
	3	0.257	26.044	0.000**
	4	0.319	27.035	0.000**
	5	0.360	29.162	0.000**
	6	0.381	31.859	0.000**

Note: ** indicate rejection of i.i.d assumption at the 5% significance levels, correspondingly. The optimal lag for each model is selected using the Schwarz information criterion

3.1 Methodology

3.1.1 Stationarity and linearity tests

In examining the linkage between country-specific (Turkey) risks vis-à-vis ERI, FRI, and PRI, the approaches of threshold cointegration, Markov-switching regression, and the frequency domain causality tests are employed to advance the studies of Kirikkaleli (2016) and Gokmenoglu, Kirikkaleli, and Eren (2019). Before this, a handful of preliminary tests are performed. To begin with, stationarity test is performed for the variables of interest. Preceding the cointegration mentioned above tests, the Zivot and Andrews (2002) stationarity approach is employed because it accounts for potential evidence of single structural break. In this case, the result of the unit root test with single structural break reveals evidence of as indicated in Table 2. The result implies that there is significant evidence of breaks in 1997Q3 and 2002Q1 for ERI. For FRI and PRI, the significant evidence of single structural break respectively exists in 1993Q2 and 2002Q1 and 1991Q2 and 1992Q4.

Additionally, the portmanteau test which known as the Brock–Dechert–Scheinkman (BDS) test by Brock et al. (1996)¹ (a detailed procedure is omitted here) is further employed to examine the nonlinear dependency property of the variables. The essence of the test is to confirm the applicability and relevance of using the nonlinear causality approach instead of the linear causality test. In evidence, the BDS test result implies that the null hypothesis that the investigated series is an iid process is rejected. Thus there is a

¹ Details and step-to-step procedure for the non-linearity test (BDS) is provided in Brock, W. A., Scheinkman, J. A., Dechert, W. D., & LeBaron, B. (1996). A test for independence based on the correlation dimension. *Econometric reviews*, 15(3), 197–235.

nonlinear dependency for all the series (see Table 3). Consequently, the investigation proceeds to apply the nonlinear cointegration techniques.

3.1.2 The cointegration tests

Considering the nonlinearity evidence from the BDS test (from Table 3 above), the threshold cointegration test with one regime shift is performed by employing Gregory-Hansen Cointegration Test (Gregory and Hansen, 1996). The evidence of nonlinearity due to potential structural changes, therefore, causing the cointegration vector to shift. Hence, these changes could be due to technological shocks or regime changes. Given that the Engle and Granger (1987) investigates the linear combination of variables (y_1 =dependent variable and y_2 =independent variable) by obtaining the residuals for the standards model from

$$y_{1t} = \mu + \alpha y_{2t} + e_t \tag{1}$$

where y_{1t} is $I(1)$, for $i=1, 2$, and e_t is $I(0)$, it is further modified by Gregory and Hansen (1996). In so doing, Gregory and Hansen (1996) adjust the intercept (μ) and/or the slope (α) such that the modified residual-based cointegration test now allows for structural changes. Then, the adjustments proposed by Gregory and Hansen (1996) present; the shift of the intercept (as in Eq. 2), the shift of the intercept with time trend (as in Eq. 3), and the shift in the intercept slope (as in Eq. 4). The modified equations are presented as

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \alpha y_{2t} + e_t \tag{2}$$

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \beta_t + \alpha y_{2t} + e_t \tag{3}$$

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \alpha_1 y_{2t} + \alpha_2 y_{2t} \varphi_{t\tau} + e_t \tag{4}$$

such that the $i=1, 2$ of μ is the respective periods before and after the regime shift, α is the slope coefficient, β_t is the time trend, and $\varphi_{t\tau}$ is the introduced dummy variable where $\varphi_{t\tau}=1$ if $t \geq [n\tau]$ or $\varphi_{t\tau}=0$ if $t < [n\tau]$ and $\tau \in (0, 1)$. Hence, the null hypothesis of no cointegration is examined by observing the residuals of the Ordinary Least Square (OLS) of Eqs. 2–4, and the single break date in each is endogenously determined. In this case, each of the variables, ERI, FRI, and PRI, is employed as a function of the other two variables in three unique models (see Table 4). The ADF test statistics are also employed to examine the significance of the break date, which is a key advantage of the Gregory and Hansen (1996) approach. The result, as presented in Table 4, shows that the null hypothesis of no cointegration is rejected in each of the models for one regime shift where the data is truncated by 15% on each side, i.e. ($\tau=(0.15, 0.85)$).

Similarly, Hatemi-J's (2008) 's approach that allows for two (regime) structural shifts during the time period is further employed. Hence, this approach offers the suitability of accounting for the two structural breaks on both the intercept and the slopes in place of one regime shift by Gregory and Hansen's (1996) approach. Although the detail and step-by-step procedure of the Hatemi-J (2008)² are not provided here, the result of the estimate and

² Details and step-to-step approach of the two regime shifts cointegration test by Hatemi-J (2008) is available in Hatemi-j, A. (2008). Tests for cointegration with two unknown regime shifts with an application to financial market integration. *Empirical Economics*, 35(3), 497–505.

Table 4 Threshold Cointegration Test

	Model 1 ERI = f(FRI, PRI)	Model 2 FRI = f(ERI, PRI)	Model 3 PRI = f(FRI, ERI)
Gregory-Hansen Cointegration Test			
ADF	-6.95 (53)**	-6.72 (52)**	-6.25(67)**
Zt	-7.03 (53)**	-6.68 (53)**	-6.39(68)**
Za	-72.49 (53)**	-64.02 (53)*	-65.37 (68)*
Asymptotic Critical Values for the Tests of Cointegration with One Regime Shift			
	1%	5%	10%
ADF	-6.45	-5.96	-5.72
Zt	-6.45	-5.96	-5.72
Za	-79.65	-6.43	-63.10
Hatemi-J Cointegration Test			
ADF*	-6.645** (0.489) (0.489)	-6.629** (0.234) (0.270)	-6.473** (0.553) (0.631)
Zt*	-15.227** (0.454) (0.496)	-6.705** (0.234) (0.270)	-6.940** (0.582) (0.582)
Za*	-178.065** (0.426) (0.496)	-79.395* (0.241) (0.270)	-42.426 (0.660) (0.702)
Asymptotic Critical Values for the Tests of Cointegration with Two Regime Shifts			
	1%	5%	10%
ADF*	-6.928	-6.458	-6.224
Zt*	-6.928	-6.458	-6.224
Za*	-99.458	-83.644	-76.806

Notes: ** and * denote statistical significance at 0.05 and 0.10 levels, respectively. The numbers in parenthesis () represent breakpoints. The Bayesian criterion determines the optimal lag for the models

especially for each model (each variable as a function of the other two) as applied above, is also provided in Table 4. The null hypothesis of no cointegration is rejected in all the cases, thus showing that there is statistical evidence of cointegration even with two regime shifts.

3.1.3 Markov switching regression and frequency domain causality

Given that the investigation reveals statistically significant evidence of cointegration in two regime shifts, the robustness is further investigated by employing the Markov-switching regression of Hamilton (1989) approach. A multivariate approach model such that the ordinary least squares (OLS) regression model is first implemented as

$$ERI_{i,t} = \beta_{0,i} + \beta_{1,i} FRI_t + \beta_{2,i} PRI_t + \varepsilon_{i,t} \quad (5)$$

where t is the quarterly periods, ε_t is the error term, and slope parameter to be estimated ($\hat{\beta}$) for each corresponding independent variable. Also, the estimation of Eq. 5 (Model A) is repeated as Eq. 6 (Model B) and Eq. 7 (Model C), where FRI and PRI are respectively the dependent variable in each case, as indicated below

$$FRI_{i,t} = \beta_{0,i} + \beta_{1,i} ERI_t + \beta_{2,i} PRI_t + \varepsilon_{i,t} \quad (6)$$

$$PRI_{i,t} = \beta_{0,i} + \beta_{1,i} ERI_t + \beta_{2,i} FRI_t + \varepsilon_{i,t} \quad (7)$$

Table 5 Markov Switching Regression

	Coefficient	Std. Error	z-Statistic	Prob
Model 1. ERI = f(FRI, PRI)				
Regime 1: high volatility				
FRI	1.050**	0.087	12.020	0.000
PRI	0.122**	0.057	2.124	0.033
C	-9.226**	4.385	-2.103	0.035
Regime 2: low volatility				
FRI	0.047	0.239	0.198	0.842
PRI	0.259*	0.137	1.882	0.059
C	2.632	4.526	0.581	0.560
Model 2. FRI = f(ERI, PRI)				
Regime 1: high volatility				
ERI	0.229**	0.042	5.339	0.000
PRI	0.090	0.065	1.378	0.168
C	22.494**	4.632	4.856	0.000
Regime 2: low volatility				
ERI	0.334**	0.048	6.907	0.000
PRI	0.187**	0.042	4.465	0.000
C	8.825**	1.621	5.441	0.000
Model 3. PRI = f(ERI, FRI)				
Regime 1: high volatility				
FRI	1.382**	0.119	11.584	0.000
ERI	-0.067	0.073	-0.911	0.362
C	21.023**	3.116	6.744	0.000
Regime 2: low volatility				
FRI	1.124**	0.128	8.772	0.000
ERI	0.295**	0.073	4.028	0.000
C	25.083	2.830	8.861	0.000

Note: ** and * denote statistical significance at 0.05 and 0.10 levels, respectively

Given that for all $\epsilon_t \sim N(0, \sigma_{st}^2)$, the switching intercept and variance of error are respectively $\beta_{0,i,rt}$ and σ_{st}^2 . Also, each of the independent variables on the dependent variable in each of the models (A, B, and C) and different regimes are respectively $\beta_{1,i,rt}$ and $\beta_{2,i,rt}$ where r_t (regime dependent) is a discrete regime variable. In addition, the latent unobserved state variable, $i = 1$ and 2 such that state one and state 2 (state of the economy) are respectively known as the *high* and *low* regimes as indicated in Table 5. The results illustrating the statistical evidence of the (2) regime shifts as shown in Table 5 offers robustness to the nonlinear cointegration evidence earlier implied in the previous cointegration tests.

Additionally, the frequency domain Granger causality test is employed to provide a robustness check to the previous estimates. Following the earlier works of Geweke (1982) and Hosoya (1991), Breitung and Candelon (2006) advanced and developed the frequency domain causality approach. The Breitung and Candelon (BC) (2006) approach provides a specific degree of variation, unlike the time-domain approach, which only offers the period of variation. Hence, the method is robust to seasonal variation and provides information

from nonlinearity and causality cycles (i.e., low or high frequency). The stepwise procedure of the BC approach is detailed below:

Given a three-dimensional vector $X_t = [ERI_t, FRI_t, PRI_t]$ of endogenous and stationary variables where time $t = 1, \dots, T$, then X_t is assumed to have a finite-order VAR representation of the form

$$\Theta(L) X_t = \varepsilon_t \tag{8}$$

where, $\Theta(L)$ is a 3 by 3 lag polynomial of order p which is presented as, $\Theta(L) = I - \Theta_1 L^1 - \dots - \Theta_p L^p$ with $L^k X_t = X_{t-k}$. The ε_t follows the white noise process with an expectation of zeros and $(\varepsilon_t \varepsilon_t') = \Sigma$, where Σ is positive and symmetric. The suitability and ease of implementing the Breitung and Candelon (2006) are also that no deterministic terms are added to the Eq. 8 above. Since Σ is positive definite and symmetric, it then offers Cholesky decomposition such as $G'G = \Sigma^{-1}$, where $G =$ lower triangular matrix and $G' =$ upper triangle matrix. Also, $E(\eta_t \eta_t') = I$ and $\eta_t = G\varepsilon_t$. In this case, the Cholesky decomposition, the MA representation of the system is given as:

$$X_t = \begin{bmatrix} ERI_t \\ FRI_t \\ PRI_t \end{bmatrix} = \Theta(L)\varepsilon_t = \begin{bmatrix} \Theta_{11}(L)\Theta_{12}(L) \\ \Theta_{21}(L)\Theta_{22}(L) \\ \Theta_{31}(L)\Theta_{32}(L) \end{bmatrix} \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \\ \varepsilon_{3t} \end{bmatrix} \tag{9}$$

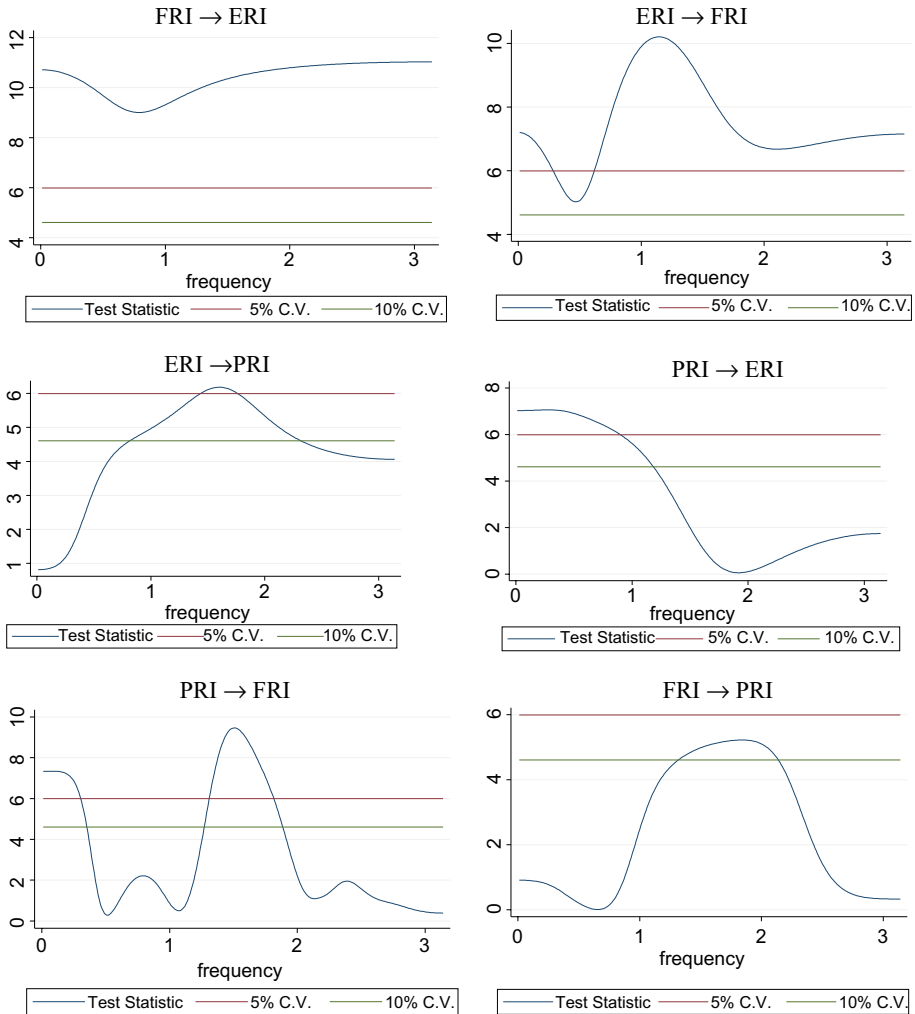
$$X_t = \begin{bmatrix} ERI_t \\ FRI_t \\ PRI_t \end{bmatrix} = \Psi(L)\eta_t = \begin{bmatrix} \Psi_{11}(L)\Psi_{12}(L) \\ \Psi_{21}(L)\Psi_{22}(L) \\ \Psi_{31}(L)\Psi_{32}(L) \end{bmatrix} \begin{bmatrix} \eta_{1t} \\ \eta_{2t} \\ \eta_{3t} \end{bmatrix} \tag{10}$$

If $\Phi(L) = \Theta(L)^{-1}$, then $\Psi(L) = \Phi(L)G^{-1}$. Consequently, the spectral density of ERI_t can be expressed as

$$f_{ERI}(\omega) = \frac{1}{2\pi} \left\{ \left| \Psi_{11}(e^{-i\omega}) \right|^2 + \left| \Psi_{12}(e^{-i\omega}) \right|^2 \right\} \tag{11}$$

The sum of two uncorrelated MA processes is represented by Eqs. 9 and 10 above. The components are driven by the past realization of ERI and the predictive power of the FRI and PRI variables. The predictive power of the FRI and PRI variables are formulated from each frequency ω in relation to the predictive component of the spectrum with the intrinsic component at that frequency. The approach of Breitung and Candelon (2006) offers a null hypothesis of no Granger causality i.e. X variable does not granger cause Y variable at frequency ω if the predictive factor of the Y variable spectrum at frequency ω is zero which is indicated by the causality tests of Geweke (1982) and Hosoya (1991) provided below

$$M_{XY}(\omega) = \ln \left[\frac{2\pi f_X(\omega)}{\left| \Psi_{11}(e^{-i\omega}) \right|^2} \right] \tag{12}$$



Note: → denotes the direction of the causality.

Fig. 2 The Frequency Domain Causality Test of Breitung and Candelon (2006) Note: → denotes the direction of the causality

$$= \ln \left[1 + \frac{|\Psi_{12}(e^{-i\omega})|^2}{|\Psi_{11}(e^{-i\omega})|^2} \right] \tag{13}$$

Also, according to Geweke (1982), the measure of causality will be zero when $|\Psi_{12}(e^{-i\omega})|^2 = 0$. However, Geweke (1982) offered a simplified linear restriction on the VAR of Eq. 1 such that

$$ERI_t = \alpha_1 ERI_{t-1} + \dots + \alpha_p ERI_{t-p} + \beta_1 FRI_{t-1} + \dots + \beta_p FRI_{t-p} + \beta_1 PRI_{t-1} + \beta_p PRI_{t-p} + \varepsilon_{1t} \tag{14}$$

where α_i and β_i ($i = 1, 2, \dots, p$) are the coefficients of the lag polynomials.

Therefore, the null hypothesis $M_{\text{ERIFRI}}(\omega) = 0$ is equivalent to the linear restriction such that,

$$H_0 : R(\omega)\beta = 0$$

where $\beta = [\beta_1, \dots, \beta_p]'$ is the vector of the coefficients of ERI, while $R(\omega)$ is as follow;

$$R(\omega) = \left[\begin{array}{cccc} \cos(\omega) & \cos(2\omega) & \dots & \cos(p\omega) \\ \sin(\omega) & \sin(2\omega) & \dots & \sin(p\omega) \end{array} \right] \quad (15)$$

The ordinary F statistic for the above VAR model of order p representation is approximately distributed as $F(2, T-2p)$ for $\omega \in (0, \pi)$, where 2 is the number of restrictions and T is the number of observations. The above test results that offer significant evidence of causality and robustness to the previous estimations are provided in Fig. 2.

4 Results and discussion

From each of the above estimations, additional information is provided supporting the statistical evidence of causal linkage of economic risk index, financial risk index, and the political risk index. The descriptive statistics (see Table 1) inform that the series ERI and FRI are not normally distributed while evidence of normal distribution is observed for the PRI. While ERI and FRI are negatively skewed, the PRI series is skewed to the right (positive skewness). The result of the Zivot-Andrew unit root test, as provided in Table 2, shows that the series are all not stationary at level. Hence, the series are all stationary after first difference (i.e. $I(0)$). There is substantial and statistical evidence that the structural break could have influenced the stationarity property with significant break dates. The economic risk index's observed break period is 2002Q1, 1993Q2, and 2002Q1 for the financial risk index, then 1991Q2 and 1992Q4 for the political risk index. The break periods 1991Q2 and 2002Q1 coincide with Turkey's general election years (Çancı and Şen, 2011).

Notably, the significant evidence of nonlinearity was preliminarily investigated and affirmed by the BDS test, thus paving the way to explore cointegration among the variables of interest. In doing this, the implemented Gregory and Hansen (1996) approach reveals evidence of cointegration in the three models adopted (see Table 4). The significant evidence of cointegration among the ERI, FRI, and PRI did not only support the results of previous studies (Kirikkaleli, 2016; Gokmenoglu, Kirikkaleli, and Eren, 2019) but in indeed suggests that the effects of risk associated with the economic, financial, and political indicators are expectedly inter-woven. Similarly, the cointegration result offered by Hatemi-J (2008), as indicated in Table 4, further affirms that there exists statistical evidence of non-linear cointegration. The proof of cointegration provided by the two approaches employed (Gregory and Hansen 1996; Hatemi-J, 2008) implies that cointegration among the ERI, FRI, and the PRI is independent of the specificity of the dependent variable.

Moreover, a robustness check is being offered by implementing the Markov-switching regression approach of Hamilton (1989) and the frequency domain causality approach by Breitung and Candelon (2006). From the Markov-switching result in Table 5; there is significant evidence of regime-switching between low volatility and high volatility in the three models employed except in the low volatility regime where the intercept of the regime shift is not significant. In the first model (i.e., ERI as a function of FRI and PRI), FRI and PRI

are observed to positively impact ERI in regime 1 (high volatility period). In contrast, only PRI exerts a positive effect on the ERI in the second regime (low volatility). In the second model (FRI as a function of ERI and PRI), both ERI and PRI exert a significant and positive impact on the FRI in the second regime (low volatility period). Simultaneously, only ERI is observed to exert a substantial and positive influence on the FRI in the first period (high volatility period). In the last model (PRI as a function of ERI and FRI), both ERI and FRI exerts a significant and positive impact on the PRI in period 2 (low volatility period). Still, only the effect of FRI on PRI is significant and positive in period 1 (high volatility period). In general, the results indicate that ERI, FRI, and the PRI are largely driven by one another, especially in the same direction. Similarly, the visual information provided by the Breitung and Candelon (2006) test in Fig. (2) implies that (red) upper lines and the (brownish) lower lines respectively represent the 5% and 10% statistically significant level.

On the other hand, the (bluish) curves indicate the statistical tests at different frequencies between the intervals of $(0, \Pi)$. Hence, statistically significant and evidence of Granger causality between the various combinations of the variables are provided except FRI to ERI. It does imply that the no Granger causality hypothesis from FRI to ERI under the specified frequency domain is not rejected. However, there exist Granger causality among the other pairs of combination as observed in Fig. 2.

5 Conclusion and policy implication

Since Schumpeter's pioneer study (1912), the linkage between economic growth and financial development has been given considerable attention by researchers, despite there is no consensus about the relationship's direction. The effect of political uncertainty on the financial and economic dynamics is one of the most intensely studied issues in the literature. However, there is a lack of empirical studies for emerging markets about the linkage between economic, political, and financial dynamics. Since investors and policymakers need to explore the link between economic, financial, and political risks, the present study aims to fill this gap in the literature for Turkey's case using the threshold cointegration Markow-switching regression and frequency domain causality tests. The present research focuses on a period 1984Q1 to 2019Q1, which involves multiple domestic and global vulnerabilities in economic, financial, and political environments. As an emerging market, Turkey faced different economic, financial and political vulnerabilities from low level and high level, to identify linkage between economic, financial and political risks became more interesting.

This study's empirical findings reveal that significant nonlinear cointegration between Economic Risk, Financial Risk and Political Risk is observed in Turkey using threshold cointegration test, which determines the structural breaks endogenously. This indicates long-run relationships between sets of time series variables, namely economic, financial and political risks. The Markow-Switching Regression outcomes show that there is a positive linkage among the risks at different volatility periods, meaning that economic, financial, and political stabilities in Turkey positively affect each other. The outcome of the frequency domain causality test of Breitung and Candelon (2006) reveals for the case of Turkey that there is feedback causality between Economic Risk, Financial Risk, and Political Risk at the different frequency levels.

As a result of newly developed econometrics techniques, it can be concluded that if governors in Turkey aim to minimize political risk, then their attention should be focused on achieving financial and economic stabilities. Moreover, to control macroeconomic dynamics, vulnerabilities in political and financial environments should be minimized. Lastly, to avoid the vulnerability in exchange rate, liquidity, and public debt, Turkey's governors consider keeping their political and economic environment stable. The study likely to open debate about the literature since the study concludes with a discussion on short-run and long-run implications for economic, political, and financial stabilises while providing policy suggestions for the policymakers in Turkey.

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Declarations

Conflict of interest Authors declare that there is no known competing financial interests or personal relationship that could have influenced the study.

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