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The Role of Financial Development in Private Sector Growth in Saudi Arabia

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Abstract

Saudi Vision 2030 (SV2030), the masterplan for the socioeconomic development of the Kingdom, places considerable emphasis on the development of the private sector in the diversification of the economy. This plan aims to have the private sector account for 65% of the total GDP of the Kingdom by 2030. As part of SV2030, the Financial Sector Development Program (FSDP), a program for the realization of targets and initiatives, was launched, aiming to enable and support financial institutions to promote the development of the private sector. Against this backdrop, we investigate the role of financial development (FD) in personal economic growth. We conduct a multivariate cointegration analysis using data from almost half a century in the extended production function framework. A key finding regarding FD is that its 1% increase leads to a 0.1% increase in the GDP of the private sector. The result does not change regardless of whether FD is measured by broad money size or bank credit to the private sector (BCP). The key policy insight from this study is that the role of financial intermediaries in private sector development should be considered together with the developments of the labor and capital markets. Two other policy insights are as follows: FD plays a positive role in developing the private sector in the long run, and the financial sector in the Kingdom should be further improved to have a more significant economic growth effect. The proposed insights are in line with the aims and objectives of the FSDP.

Keywords: Saudi Arabia, private sector, financial development, multivariate cointegration.

1. Introduction

There is not much need, except for in oil-exporting developing countries, to justify the importance of the private sector in developing national economies. Governments play a leading role in socioeconomic development in such economies, mainly through fiscal policies. However, the literature shows that the private sector should primarily drive long-term sustainable growth. Since 2016, with the announcement of Saudi Vision 2030 (SV2030)—a strategic roadmap for the development of Saudi Arabia—the Kingdom has embarked on the path to diversify its economy beyond hydrocarbons. One of the main goals of SV2030 is to increase the private sector's contribution[1] to GDP from 40% in 2016 to 65% in 2030.

There are various sources of economic development, including private sector growth. The financial intermediation and development of the financial sector are extensively considered in this respect, as the financial sector plays the essential role of intermediary, acting as a capital allocation mechanism. Notably, SV2030 considers financial development (FD) among the key enablers for the future development of the Kingdom. Therefore, one of the vision realization programs, the Financial Sector Development Program (FSDP), is dedicated to the developmental aspects of Saudi financial markets. This program aims to establish a diversified and effective financial services sector to boost the development of the Kingdom's economy; diversify income sources; and stimulate savings, finance, and investment. To achieve this aim, the program has three main objectives: (i) enabling financial institutions to promote the development of the private sector, (ii) ensuring advanced capital market formation, and (iii) promoting and enabling financial planning (retirement, savings, etc.) without impeding the strategic objectives intended to maintain the stability of the financial services sector. Another point worth considering is that the abovementioned three objectives are also considered among SV2030 objectives. Moreover, the FSDP indirectly contributes to more than 20 SV2030 objectives, many of which are designed to support private sector development (FSDP 2021).

The abovementioned background highlights the importance of FD for the future development of the private sector in the Kingdom. Therefore, understanding how and to what extent FD can contribute to supporting private sector growth is crucial for policy-makers, as it is one of the enablers of nonoil sector development via capital transmission. This paper aims to assess the role of FD in developing the private sector in Saudi Arabia, the world's largest oil exporter and a member of the G20, and to derive insights that might be useful for policy-making.

We quantify the impact of FD on private economic development in the extended production function framework using a multivariate cointegration method for the period 1970-2018. We find that FD, alongside capital and labor, exerts a positive effect on private sector development. The elasticity is approximately 0.1 regardless of whether it is measured by broad money size or bank credit to the private sector (BCP).

The key policy insight of this work is that the private sector's development, including the role of financial intermediaries in this development, should consider the actions of the labor and capital markets in Saudi Arabia. Moreover, FD can play a positive role in the development of the private sector in the long run. Furthermore, the financial sector in the Kingdom should be further developed to exert a larger economic growth effect. In other words, there is still a large role for FD to play in developing the private sector. A need for the further improvement of the financial sector is also highlighted in the FSDP.

There are a limited number of studies on the relationship between FD and economic growth in Saudi Arabia, sometimes with contrasting findings regarding both direction and magnitude. Hence, we believe that this paper has value for the literature, as it has the following merits. First, the empirical analysis of this work lies in a theoretically justified framework rather than in considering variables of interest in an ad hoc manner. Second, the findings and, thus, policy insights of this study are derived from a multivariate cointegration (long-run) analysis, which has obvious advantages over a single equation-based analysis (see Section 6). Third, to obtain robust results and make policy recommendations that are more informed, this study considers two measures of FD. Fourth, unlike previous studies, the study period includes the domestic energy market and fiscal reforms alongside other transformational changes and oil price decreases.

The rest of the paper is structured as follows. Section 2 provides a background of the Saudi Arabian financial sector, while Section 3 surveys the related literature. Section 4 describes the theoretical framework of the study. Section 5 presents the data, and Section 6 provides empirical specifications and the econometric approach used. Section 7 documents the results of the empirical analysis, while Section 8 discusses these results. Finally, Section 8 concludes the paper with a few policy insights.

2. Background

2.1. Banking Sector Development

The key benefits of the banking sector are that it mobilizes financial resources to be used in the most efficient manner possible. The Saudi Arabian banking structure is atypical of Western models because it is based on Islamic fundamentals. The use of usury is prohibited. Thus, the distribution of risk is shared between the lender and borrower according to Sharia law. Two significant forms of financing are allowed: *musharaka* financing, in which the borrower and lender agree to a joint venture to split the risk of loss and the fruits of profits in the proportion of the amount invested, and *mudaraba* financing, which is formed when one party finances the capital requirements for a project and a second party offers its entrepreneurial and managerial skills to run a project or business. Therefore, Saudi banks act more as financiers than as traditional banks. Thus, it is expected that the extent and depth of intermediation might be depressed due to fewer risk-taking measures imposed by the structure of Islamic laws concerning loan issuance. Therefore, the role of intermediation may be operating below that of its Western counterparts.

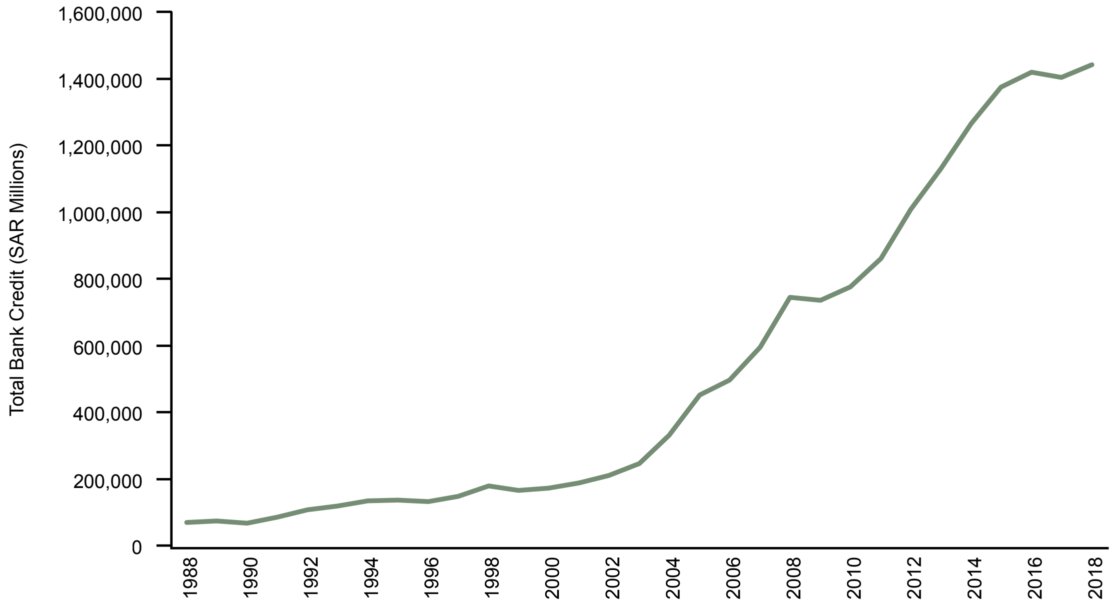
Only six banks operated in the Kingdom in 1952, regulated by the Saudi Central Bank, formerly known as the Saudi Arabian Monetary Authority (SAMA). While the SAMA existed then, it had little control over the country's financial system. The government began to control and regulate the financial system due to the near failure of Bank of Riyadh in 1964. The Banking Control Code was issued in 1966, which standardized the operations of the banking sector. As a result, bank licenses grew in number in the 1960s and 1970s. By 1975, fourteen banks, in addition to specialized banks, were operating in the country.

During the 1985-1995 developmental plans of the Saudi economy, the emphasis was placed on the Kingdom's financial institutions and included prompting national financial institutions to invest in private funds, extending credits to production projects rather than to import trade, and encouraging more joint-stock companies, among others. However, during this era, the decline in government revenues due to the collapse of oil prices sent the banking sector into distress. The lack of risk management of portfolios in banks was apparent when government austerity measures took effect.

Banks provide an essential measure of FD: BCP. This indicator provides a strong representation of the role of intermediation, in which banks channel monies from savers to investors. Levine, Loyaza, and Beck (2000) highlighted that BCP is a good indicator to represent the role of intermediation, as it serves a functional role in capital markets and excludes credit offered to the public sector.

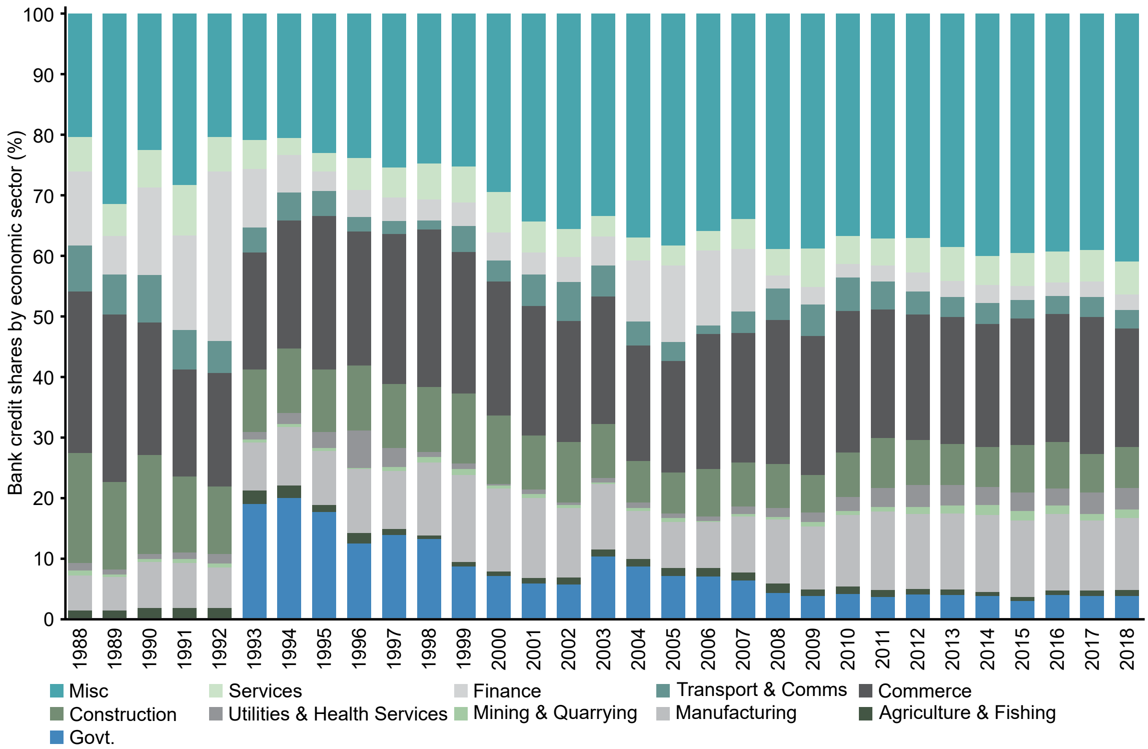
Figure 1 depicts total bank credit by economic activity for Saudi Arabia from 1988 to 2018. There was steady and slow growth until 2003 followed by a large increase in the growth rate. Overall, total bank credit grew by approximately 10.4% per year. Examining the shares of total bank credit reveals which economic sectors have appropriated resources channeled through the banking sector. Figure 2 displays the BCP and government by their share of the total share of bank credit from 1988 to 2018.

Figure 1. Total Bank Credit by Economic Activity.



Source: SAMA (2018).

Figure 2. Bank Credit Shares by Economic Sector.



Source: SAMA (2018).

2. Background

Table 1. Bank Credit by Economic Activity.

	Government	Agriculture and Fishing	Manufacturing	Mining and Quarrying	Utilities and Health Services	Construction	Commerce	Transport and Communications	Finance	Services	Misc
1990		1.8	7.7	0.4	1.0	16.4	21.9	8.0	14.2	6.4	22.3
2000	7.2	0.7	13.7	0.4	0.4	11.2	22.1	3.6	4.7	6.7	29.4
Delta		-1.1	6.0	0.0	0.6	5.2	0.2	4.4	9.5	0.3	7.0
2010	4.2	1.3	11.6	0.8	2.5	7.2	23.4	5.5	2.3	4.6	36.7
Delta		0.6	-2.1	0.3	2.1	-4.0	1.2	2.0	-2.4	-2.1	7.3
2018	3.7	1.0	12.0	1.3	3.6	6.7	19.6	3.0	2.6	5.5	40.9
Delta		-0.3	0.4	0.6	1.1	-0.4	-3.8	-2.5	0.3	0.9	4.3

Values are expressed as percent (%).

Sources: SAMA (2018).

The most considerable fluctuations in bank credit shares occurred between 1990 and 2000, and bank credit increased the most during this period for the manufacturing sector, at 6%, and for the miscellaneous sector, at 7%. In contrast, the levels of construction and finance decreased by 5.2% and 9.5%, respectively. By 2018, the largest share of banking credit was channeled to the miscellaneous sector, which saw a significant rise between 2000 and 2010, followed by the commerce sector, at 19.6%. Other important economic sectors had less credit flow shares in 2018 compared to those in 2010, with a decrease for the commerce, transport and communication, and agriculture sectors. The underpinnings of these findings shed light on the allocations made by credit institutions to different economic sectors. Credits provided by sectors with high levels of imports and less productivity imply that credit growth may not reap economic benefits, as it would in other sectors. The credit provided by banks provides capital resources that sectors can access to expand their operations and finance new projects and ventures.

2.2. Monetary Base

Saudi Arabia has pursued a policy of fixed exchange rates since 1987. More specifically, the riyal is pegged to the US dollar with a fixed rate of 3.75. Theoretically, the impossible trinity, known as the Modell-Fleming condition, articulates that having a fixed exchange rate regime implies either the loss of monetary policy independence or the controlling of international capital flows. Saudi Arabia does not control international capital flows, and the implementation of this fixed exchange rate policy has contributed significantly to economic development, especially for the diversification and expansion of the nonoil sector, as previous studies have concluded (see Alkhareif, Barnett, and Qualls. 2017; Razek and McQuinn 2021). To maintain the stability of the rate of 3.75 riyals per US dollar, the interest rate policy of Saudi Central Bank follows the interest rate policy of the US Federal Reserve, as interest rate parity is one of the requirements of the fixed exchange rate system (see Al-Gahtani 2015, among others). This approach has two main policy implications: the interest rate cannot be used as a tool in monetary policy, and excess money supply, when demand is at a higher level, cannot be generated in the economy. It turns out that Saudi Central Bank should not create a surplus money supply in the economy to boost economic activities because this excess money would place pressure on the domestic and foreign prices of the riyal, that is, the interest and exchange rates, which could result in direct and/or indirect problems for the stability of the fixed exchange rate. Since the inefficiencies of monetary policy in stimulating economic growth in fixed exchange rate regimes are well documented in the theoretical and empirical literature, we do not discuss them here. Moreover, we do not discuss how fiscal policy is generally effective in boosting economic growth in countries with a fixed exchange rate regime or how it is implemented in Saudi Arabia. Our finding that expanding the level of broad money and bank loans supports economic development in the private sector, among other things, implies the following. Saudi Central Bank has successfully provided the demanded amount of money, and thus, it has created the required liquidity and thus stimulated the development of the private sector without creating a severe problem regarding the stability of the fixed exchange rate of the riyal. In other words, the monetary authority has met the demand for money by the economic agents needed for their activities. Nevertheless, this authority has not generated excess money, which can result in pressure on the pegged exchange rate.

3. Literature Review

The nexus between FD and economic growth has been investigated comprehensively in the literature. However, it is well known that financial systems play the essential role of intermediaries. This intermediary status functions as a mechanism that matches a borrower with a lender and savings to investments. FD encompasses the important role of capital mobilization in the most efficient uses of such financial systems, which begs the question of to what extent FD contributes to economic growth. Nevertheless, the quality and fluidity of intermediation are contingent on the ability and efficiency of financial institutions. Theory and literature impose a set of indicators that identify the effects of this intermediation on economic growth, such as bank credit, broad money aggregates, and productivity in the financial sector.

There are two main opposing views regarding how FD contributes to economic growth. The first view is the 'supply-leading' argument, which states that FD promotes economic growth by behaving as an input to economic growth. These views have been put forward by Schumpeter (1943), Patrick (1966), McKinnon (1973), and Shaw (1983). Other advocates suggest that FD incentivizes savings, facilitating capital accumulation and leading to improved investments and growth Galbis (1977). Fry (1978), Goldsmith (1969), Greenwood and Jovanovic (1990), Thakor (1996), and Hicks (1969). The second view, the 'demand-following' view, argues that FD depends on economic development. This suggests that as an economy develops, demand for financial services increases (Robinson 1952). As a result, financial institutions have begun to create additional financial instruments, leading to innovations in the financial sector. Al-Yousef (2002) and Ang and McKibbin (2007) confirm this hypothesis empirically. While theory and empirical findings are in conflict regarding the direction of causality between FD and economic growth, empirical evidence suggests that FD influences economic growth and does not exhibit simultaneity (Demirgiic-Kunt and Levine 2008; Levine 2005).

The literature on the effect of FD on economic growth dates to the nineteenth century, where Bagehot (1873) suggested that financial systems played a crucial role in the Industrial Revolution in England. Hicks (1969) found that the allocation of capital facilitated by the development of financial markets was a key enabler for industrial progress. Several financial indicators were found to be positively correlated with economic growth. Goldsmith (1969) found that the size of the financial system was positively related to economic activity, using 35 countries from 1860 to 1963. McKinnon (1973) and Shaw (1973) found that financial markets played a pivotal role in economic activity. King and Levine (1993) found a positive relationship between bank credit and economic growth. When examining cross-sectional data from 80 countries, Levine and Zervos (1998), Levine (1998), and Beck and Levine (2003) found similar results. While these studies showed how FD facilitates economic growth, they did not explicitly deal with the issues related to causality. Eschenback (2004) found that the causal link varied across countries regarding direction and variable selection. An examination of the literature reveals that opinions are divided regarding the direction of causality between FD and economic growth. Several studies have adopted instrumental variables as a remedy to deal with endogeneity. A proxy for FD has been used to impose the exogenous variation in FD. Moreover, the measure of legal origin of La Porta et al. (1998) has been widely adopted. Using the instrument of FD with generalized methods of moments (GMM) estimators, Levine (1998, 1999) found that FD contributes to economic growth, capital accumulation, and productivity.

As with the eminent role of intermediation by the banking sector, another channel of intermediation is through financial markets. Stock market prices are considered a leading indicator of economic performance (Fama 1981, 1990). In financial markets, higher stock prices trigger households and firms to participate in transactions. Moreover, higher stock prices suggest greater confidence and lower uncertainty when assessing the state of an economy. The theory that relates the stock market to the economy of Modigliani (1971) proposes that stock market performance increases household wealth and income, in turn adjusting their consumption levels.

Moreover, stock prices influence publicly traded firms' balance sheets. When stock prices rise, higher-valued firms can obtain larger credit allocations for investments and operations, leading to an increased economic development level (Gertler and Bernanke 1989). While the prescribed evidence and literature on the relationship between financial market performance and economic growth are intuitive (Barro 1990; Fama 1981; Humpe and Mamillan 2005; Mauro 2003; Stewart 1990;), other studies are not (Binswanger 2000, 2004).

Some studies have stressed that FD exhibits heterogeneous effects on economic growth, contingent on whether the countries in question are developing or industrial. Rioja and Valev (2004a) found that FD positively impacts economic growth via capital accumulation in developing countries, whereas in developed economies, the effects are transmitted through productivity growth. In a study using 119 countries by Deidda and Fattouh (2012), they found that FD exerts a significant and positive influence on richer economies. However, this influence is nonsignificant for poorer economies. In contrast, 71 countries show significantly positive and larger effects for poorer economies than for richer ones (Huang and Lin 2009). Other studies found weaker relationships when considering more recent data, suggesting the vanishing effects of FD (Rousseau and Wachtel 2011). Arguing that financial deepening need not be isolated from the adequate reform for the continuous improvement of financial sector regulations, this finding suggests that the relationship between growth and FD is more complex than shown in the simplified empirical results. This realization could explain the growing amount of research proposing that the finance-growth nexus relationship is nonlinear. The above authors showed that when BCP reaches the threshold of 80-120% of GDP, the extent of financial depth is exhausted and begins to exert a negative effect on the economy. Moreover, Samargandi, Fidrmuc, and Ghosh (2015) exploited data on 52 middle-income countries and found that FD and economic growth exhibit an inverted U-shaped relationship. This is also seen in the work of Rioja and Valev (2004b), who suggested that the positive effect is observed only once economies reach a specific threshold.

This paper aims to investigate the impact of FD on private sector growth in Saudi Arabia. Historically, oil revenues were the main driver of socioeconomic development in the Kingdom. Since the price collapse in late 2014, the government has promoted many initiatives spearheaded by Vision 2030 to diversify the Kingdom's economy. As a vital vision realization programs, according to the FSDP Charter (FSDPC), the FD of the country is a priority. The FSDPC is a delivery plan for financial sector development to create a growing financial sector, which is crucial to unlocking the potential of the Saudi economy. This charter aims to increase the values of the essential indicators of FD, such as the amount of BCP, and to increase the level of stock market capitalization to provide adequate funding to meet the aspirations of the Kingdom's Vision 2030 (FSDPC 2018).

3. Literature Review

Several empirical studies have been performed to examine the relationship between FD and economic growth for Saudi Arabia. However, as with the overall literature, the literature investigating the impact of FD and that includes Saudi Arabia is divided. Darrat (1999) found that the finance-growth nexus exerts both a supply-leading and demand-following effect on Saudi Arabia. Xu (200) used multivariate VAR to investigate FD's effects on GDP in 41 countries, including Saudi Arabia, finding that the long-term elasticities are negative. Al-Tamimi (2002) did not find evidence of either the significance of the role of FD for economic growth for Saudi Arabia or the effect of economic growth on FD. The above author explained that these results are due to immature financial systems and deficient financial instruments.

Al-Yousif (2002) used a Granger causality test for 30 developing countries, including Saudi Arabia, finding that FD, measured by M2/nominal GDP (King and Levine 1993), does not Granger-cause GDP growth. Boulia and Trabelsi (2015) found no evidence of a causal relationship either way with all the indicators of FD for the Saudi economy and stated that these results may indicate the absence of a relationship between FD and economic growth. Similarly, Ibrahim (2013) examined domestic BCP by exploiting data from 1989 to 2008. Using fully modified ordinary least squares (FMOLS) estimates, domestic BCP was shown to have a significant and positive impact on economic growth in the long term but a nonsignificant and negative impact on economic growth in the short term. Al-Awad and Harb (2014) found that a one-way causality runs from FD to economic growth. Only Saudi Arabia, of a sample group of oil-producing countries, has achieved cointegration. Samargandi et al. (2014) found that the impact of FD, measured as a single composite indicator through principal component analysis (PCA) of three indicators—M2/GDP, M3/GDP, and BCP—has a positive impact on the growth of the Saudi nonoil sector. In contrast, for the oil sector, this effect is negative or nonsignificant. The above authors argue that the relationship between FD and growth may fundamentally differ between resource-dependent and resource-independent economies. Furthermore, Alghfais (2016) conclude that financial sector development has a positive and significant impact, proxied by six indicators in PCA, on the private economy using an autoregressive distributed lag model.

4. Theoretical Framework

As discussed previously, this study aims to primarily analyze the effect of Saudi financial sector development on economic growth. Many theoretical and empirical studies have discouraged the use of a bivariate framework, as they argue that it can lead to misleading conclusions, such as false Granger causality results (e.g., Caporale and Pittis 1997; Lutkepohl 1982; Odhiambo 2009; Triacca 1998). The literature also indicates that the relationship's magnitude and direction may differ once an additional variable is introduced.

We base our study on the extended production function framework as described below. Note that the augmentation of the production function framework with the variable of interest is widely used in theoretical and empirical studies (e.g., Feder 1982 for exports; Grossman 1988 and Alexiou 2009 for government spending; Hasanov et al. 2021b for government spending in Saudi Arabia).

The standard neoclassical Cobb-Douglas production function in the case of constant returns to scale can be written as follows:

$$Y = AK^{a}L^{1-a} \tag{1}$$

where Y, K, and L are output, capital, and labor, respectively; a and (1-a) are the capital and labor elasticities of the output, respectively. A denotes total factor productivity (TFP).

The literature suggests that the impact of FD on output (or per capita output) can operate through TFP (see Bolbol et al. 2005). We highlight the role of FD in investment efficiency, capital allocation, and financial deepening, particularly as an enabling condition. Hence, following Dasgupta, Keller, and Srinivasan (2002) and Badeep et al. (2016), we consider TFP to represent investment efficiency. Furthermore, following Badeep et al. (2016), Nawaz, Lahiani, and Roubaud (2019), and Krinichansky and Sergi (2019), we consider TFP a function of FD and other factors (X_n).

$$A = TFP = f(FD_t, X_o) \tag{2}$$

Equation (1) can be replaced by Equation (2) as follows:

$$Y = (FD, X_o)K^{a}L^{1-a} \tag{3}$$

It is reasonable to assume that FD changes faster than do the other factors of TFP, such as institutional development, technological progress, and the establishment of an efficient business environment, for which a very long timeframe is needed to achieve changes. Hence, we treat FD as changing over time, such as capital and labor, while other factors are considered time invariant in the medium to long run. Resultantly, Equation (3) can be written as follows:

$$Y = X_o K^{a}L^{1-a}FD^{b} \tag{3A}$$

The log-linearization of Equation (3A) in the time-series context, denoted by subscript t, yields the following expression:

$$y_t = c + a_1 k_t + a_2 l_t + bfd_t + e_t \tag{3B}$$

where c = exp (X_o) and exp is the exponent function. Lower-case letters indicate the natural logarithmic expression of the upper-case letters, $a_1 = a$ and $a_2 = 1 - a$, b is the elasticity of output with respect to FD, and e_t is the error term.

5. Data

In line with the above extended production function, we collect annual time-series values of the variables for the period 1970-2018. To check different specifications for robustness, we consider two widely used measures of FD: bank claims in the private sector and the broad money supply share of GDP.[2] Table 2 documents the descriptions and sources of the variables.

Table 2. Variables and their descriptions.

Variable	Notation	Description	Source
Nonoil private sector value added	Y	The variable is defined as GDP excluding the mining and quarrying sector and net taxes as well as the value added created by the producers of the government services and other goods and services. This variable is measured in a million SAR at 2010 prices. For simplicity, we refer to this variable as nonoil private GDP throughout the paper.	GaStat (2018)
Broad-money-to-GDP ratio	M2	This is the ratio of M2 monetary aggregate to GDP. M2 is the sum of M1 (narrow money aggregate) and time and saving deposits, in million SAR. The time series of the variable is retrieved from the SAMA (2018). GDP is the sum of value added produced in all sectors of the KSA economy in real terms of million SAR at 2010 constant prices. The time series of the variable is collected from GaStat (2018).[3]	Calculated
Bank claims in the private sector	BC	These are the loans that banks provide to the private sector, measured in a million SAR at 2010 prices. Information on bank claims in the private sector in nominal million SAR is collected from the SAMA (2018). This variable is deflated by the nonoil GDP deflator to convert the values into real terms. The latter variable is the percentage ratio of nominal nonoil GDP to real nonoil GDP at 2010 prices. Both nominal and real nonoil GDP series are collected from GaStat (2018).	Calculated
Nonoil private employment	L	This is the total number of people employed in the nonoil sector, excluding those in the government sector, measured in thousands of people. Information on nonoil employment is collected from the CIEC (2018) for 1970-1998 and GaStat (2018) for 1999-2018. Information on government sector employment is collected from the SAMA (2018).	Calculated
Nonoil private sector capital stock	K	This is the adjusted capital stock in the nonoil private sector in million SAR at 2010 prices and is constructed using the perpetual inventory method, as expressed below (Collins, Bosworth and Rodrik 1996; Caselli 2005): $$K_t = (1-\kappa)K_{t-1} + I_t$$ where K_t represents nonoil private sector capital stock at time t, κ is the depreciation rate, and I_t is the nonoil private sector investment at time t. Nonoil private sector investment in million SAR at 2010 prices is constructed as follows. Nominal nonoil private sector investment in million SAR (the values for 1970-1995 are from the SAMA (2016) and for 1996-2018 are from the SAMA (2018)) deflated by investment deflator, 2010=100. The investment deflator is the percentage ratio of nominal gross fixed capital formation to real gross fixed capital formation at 2010 prices, collected from GaStat (2018) and GaStat annual yearbooks, respectively. We assume a 5% depreciation rate and the initial level of the capital stock to be 1.5 times that of nonoil private GDP in 1970. Finally, we exclude bank claims in the private sector (BC) from the resulting capital series to avoid double accounting in the empirical analysis, following Francois and Keinsley (2019) and Herzer and Morrissey (2013), inter alia.	Calculated

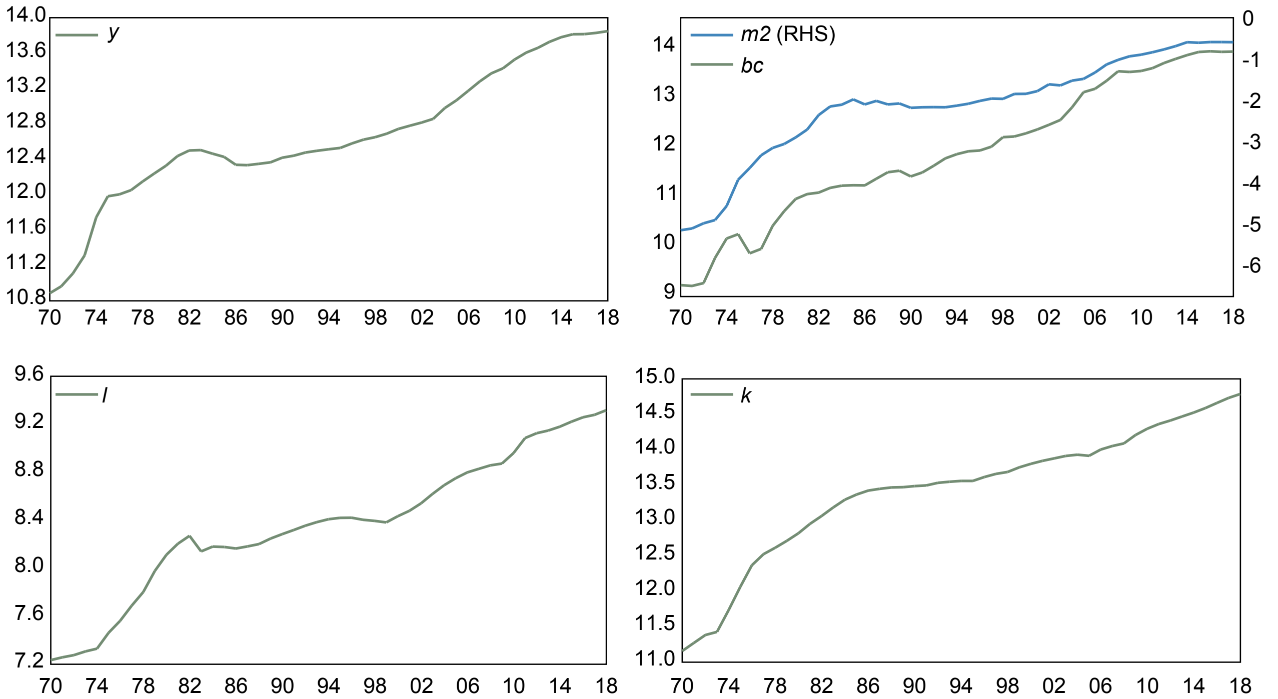
Notes: The GaStat is the General Authority for Statistics of Saudi Arabia, the SAMA is Saudi Central Bank (the former Saudi Arabian Monetary Authority), and the CIEC is the CIEC database.

We use the natural logarithm expression of the variables (denoted by lowercase letters) and their differences (indicated by Δ) in the empirical analysis, as illustrated in Figure 3.

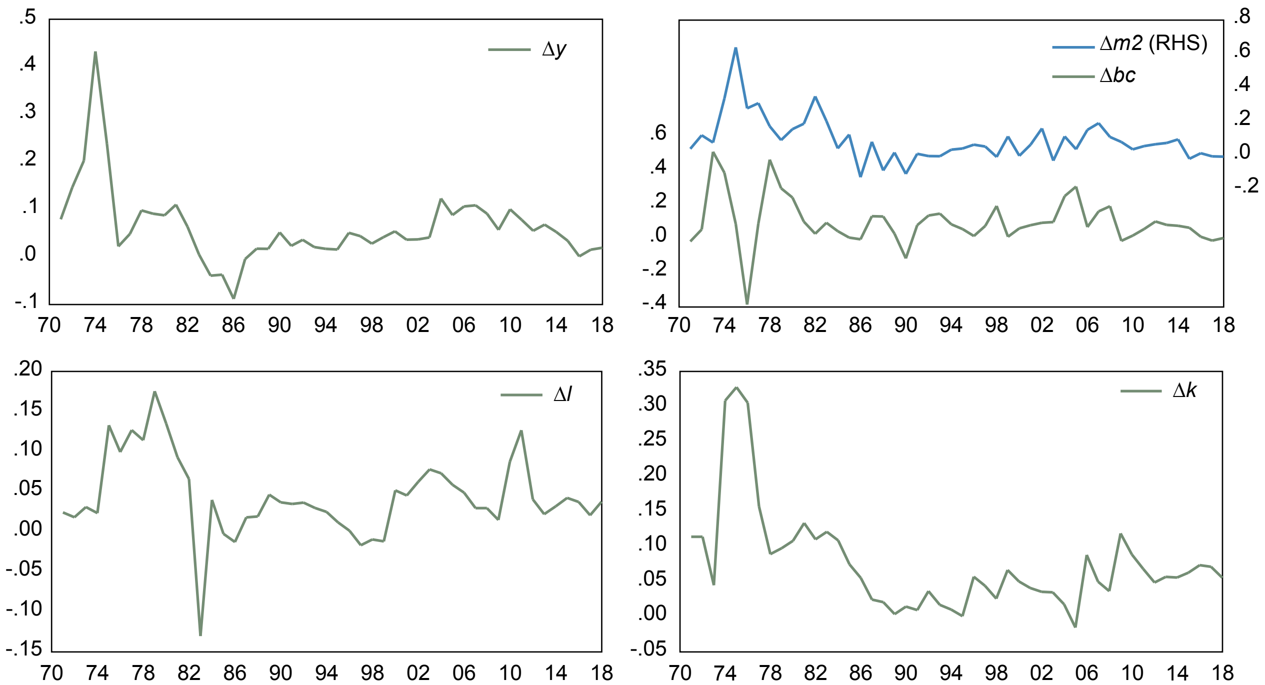
Figure 3 highlights the time development of the variables of interest (*M2* and *BC*) and control variables (*K* and *L*) alongside our dependent variable—private sector value added (*Y*).

Figure 3. Logarithmic levels and growth rates of the variables.

Panel A. Log levels of the variables



Panel B. Growth rates of the variables



5. Data

Table 3. Descriptive statistics of the variables, 1970-2018.

	y	*m2*	*bc*	*l*	*k*
Mean	12.65	-2.02	11.91	8.37	13.43
Median	12.51	-1.89	11.88	8.38	13.56
Maximum	13.85	-0.45	13.96	9.32	14.79
Minimum	10.88	-5.03	9.21	7.23	11.15
Std. Dev.	0.74	1.25	1.40	0.58	0.93

Since 1970, *Y* has grown 18-fold, averaging 6.72% per year. Nonoil GDP has grown in terms of its contribution to total GDP from 11% of GDP in 1970 to approximately 40% in 2018 (SAMA 2019). Private sector value added has become a central focus of policy-makers to improve its contribution to GDP, further aligning it with Vision 2030 (Vision 2030 2017). *BC* and *M2*, the measures for FD, have developed to some degree as a response to real activity growth. *BC*, in real terms, grew very quickly in 1970, averaging 23% per year from 1970-1980 and 8.5% from 1980-2018. *M2* followed similar rapid growth during 1970-1980, catching up to economic activity and growing 8.4% per year from 1980-2018. Table 3 presents the descriptive statistics of the natural logarithmic transformations of *Y*, *M2*, *BC*, *L,* and *K* for the entire period.

Notably, the distribution of the log level of the variables is quite similar regarding first and second moments. The highest standard deviations are recorded for the FD measures compared to the control variables.

6. Empirical Specifications and the Econometric Approach

We follow the theoretical framework discussed in Section 4 in the empirical analysis. To obtain robust results and suggest informed policy insights, we estimate two specifications for the effects on private sector GDP effects of FD, as given below.

$$y_t = \beta_0 + \beta_1 m2_t + \beta_2 \ l_t + \beta_3 k_t + \beta_4 t_t + \varepsilon_t \qquad (4)$$

$$y_t = \beta'_0 + \beta'_1 bc_t + \beta'_2 \ l_t + \beta'_3 k_t + \beta'_4 t_t + v_t \qquad (5)$$

where *y*, *l*, and *k* are the natural logarithm expressions of *Y*, *L*, and *K*, respectively; *m*2 and *bc* are the logarithm expressions of *M2* and *BC*, respectively; and $\beta_0, \beta_1, \beta_2, \beta_3, \beta_4$ and $\beta'_0, \beta'_1, \beta'_2, \beta'_3, \beta'_4$ are coefficients to be econometrically estimated. ε and v are the error terms. *t* denotes time.

The econometric analysis of this research covers unit root and cointegration tests, as well as estimations of the long-run coefficients. We employ an augmented Dickey-Fuller (ADF) test (Dickey and Fuller 1979) and a Philips-Perron (PP) test (Phillips and Perron 1988) for robustness, as they are widely used in empirical analyses.

Johansen's multivariate cointegration method is considered the primary method in this work. Because a system-based cointegration method has numerous advantages over a single equation-based cointegration method, as the literature discusses, we try to present an overview of them below. First, n-1 cointegrating relations can be established among n variables. However, single equation and residual-based methods can detect only one relationship (Badinger 2004; Enders 2015; Johansen 1988; Johansen 1991). The point here is that the incorrect identification of the number of cointegrating relations that established by the variables (e.g., considering one cointegration of interest and ignoring others) leads to, at best, information loss. In the worst case, this situation can cause biased and inefficient estimates, as Juselius (1992, 1995) and Phillips (1991), among others, have discussed.[4] Dibooglu and Enders (1995), among others, noted that a single equation-based method may not detect even a single cointegrated relation if the variables under consideration establish more than one relation. Moreover, Kugler and Lenz (1993) discussed that single-equation cointegration methods break down if the regressors are cointegrated. Furthermore, system methods are not subject to the normalization problem and provide a convenient framework through which to test the theoretical hypotheses, unlike single equation methods (e.g., Enders 2015; Juselius 2006). Another problem with the single equation methods that should cause care to be taken among researchers is that they do not provide information about the weak exogeneity of the regressors by default, and thus, a weak exogeneity test should also be conducted, as was done by Narayan (2004) and Narayan and Russel (2005), among others. The weak exogeneity of the regressors is essential if their contemporaneous values are to be included in the short-run model of the variable of interest. Otherwise, the estimates are inefficient (e.g., Ericsson and MacKinnon 2002; Juselius 2006). The above discussion emphasizes that a single equation cointegration method can be used if the variables under consideration establish only one long-run relation. If these variables establish more than one cointegrating relation but a single equation model is preferred, then a conditional modeling exercise has to be conducted. In a nutshell, such an exercise involves identifying other cointegrating relations in the system framework, such as the Johansen method, in addition to the relation of interest and including all of relations in the short-run equation, i.e., the

ECM of the variable of interest, while accounting for weak exogeneity (see Boswijk 1995; Ericsson 1995 inter alia).

Once we determine that the variables are cointegrating, we estimate the long-run equation and speed of adjustment (SoA) coefficients through the vector equilibrium correction (VEC) model framework. Additionally, we test for the significance of various restrictions to identify long-run relationships. Appendix 1 provides the details of the econometric methods employed.

7. Empirical Results

7.1. Unit Root Test Results

Table 4 reports the results of the ADF and PP tests.

For each variable, a detailed discussion of the results of both tests can be found in Appendix 2A. The Appendix discusses that it is straightforward to conclude that the measures of FD, that is, *bc* and *m2,* follow the I(1) process, meaning that their log levels contain a unit root; however, the first differences of the log levels are stationary. The discussion in the Appendix concludes that *y*, which is our dependent variable, and *l* and *k* also follow the I(1) process. Since the order of integration of the variables is not greater than one, we can use Johansen's reduced rank method for the cointegration analysis in the next step.

7.2. Results of the Long-Run Analysis

We find three cointegrating relationships among the variables in Equation (4), i.e., private GDP, broad-money-to-GDP ratio, private capital, private labor, and time trend. The time trend factor represents the known and unknown variables that affect our dependent variable over time but that are not measured and included in the equation (see Kim and Heshmati 2019; Nelson and Kang 1984 inter alia). Appendix 2B discusses the results of the estimation, stability, and postestimation tests for the VAR model while also discussing the cointegration test results performed on the transformed VECM. Here, we report only the identified long-run relationship between FD and private GDP, relegating

Table 4. URT test results.

Variable	ADF test					PP test			
	Test value	*C*	*t*	*None*	k	Test value	*C*	*t*	*None*
y	-3.81***		x		1	-2.69		x	
m2	-3.07		x		1	-2.21		x	
bc	-3.05		x		2	-1.64		x	
l	-3.32*		x		2	-1.13	x		
k	-3.45*		x		1	-2.96		x	
Δy	-2.32**			x	0	-2.07**			x
Δm2	-3.63***	x			0	-3.70***	x		
Δbc	-6.93***	x			1	-4.72***	x		
Δl	-3.84***	x			0	-3.91***	x		
Δk	-2.68*	x			0	-1.64*			x

Notes: ADF and PP denote the augmented Dickey-Fuller and Phillips-Perron tests. The maximum lag order is set to three, and the optimal lag order (***k***) is selected based on the Schwarz criterion in the tests; ***, **, and * indicate the rejection of the null hypotheses of the unit root at 1%, 5%, and 10% levels, respectively. The critical values for the tests are taken from MacKinnon (1996). None denotes that neither the intercept nor the trend is included in the test equation. The final UR test equation can include one of three factors—intercept (***C***), intercept, and trend (***t***)—or none of them (***None***). x indicates that the corresponding option is selected in the final UR test equation.

the other two relationships to Appendix 2.B in case readers are interested. The Appendix also discusses the details of the imposed restrictions to identify long-run equations for private GDP, labor, and capital.

$$y_t = 1.96 + 0.10\ m2_t + 0.52\ l_t + 0.48\ k_t + 0.003t_t + \varepsilon_t \quad (4')$$

The number of cointegrated relations does not change if we measure FD with bank claims in the private sector instead of the broad-money-to-GDP ratio, that is, using Equation (5). This might indicate that the results of the cointegration analysis are robust and worthy of consideration. The results for specifying a VAR model, its stability and postestimations tests and the cointegration test results from the transformed VECM are reported in Appendix 2C. For reader convenience, only the identified long-run equation of private GDP is reported here. The other two identified relationships for private sector labor and capital are provided in Appendix 2C, along with the statistical and theoretical details of the identification of these long-run relationships.

$$y_t = 0.69 + 0.10\ bc_t + 0.52\ l_t + 0.48\ k_t + v_t \quad (5')$$

8. Discussion of Empirical Results

The unit root tests show that the natural logarithmic expressions of *Y, M2, BC, L*, and *K* are nonstationary but that their first differences are stationary; i.e., they are all I(1) processes. The nonstationarity of the variables has two main implications: first, when shocks act on the variables, they can have permanent effects, and the variables do not return to their previous mean, and second, the stationarity of the variables implies that they are mean reverting, which means that the shocks acting on them can produce only temporary, not permanent, changes.

We conclude that there are three cointegrated relationships among the variables being identified for private GDP, private labor, and private capital stock. This finding means that the mentioned variables and their regressors move together in the long run, as they share a common trend/path. The presence of cointegration implies that the identified relationships are theoretically or empirically interpretable and hence not spurious and that the numerical values (i.e., estimated coefficients) of these relationships are valid for discussion and/or policy considerations.

According to Equation (4'), a **1%** increase in the level of FD, represented by a broad money-to-GDP ratio, translates into a **0.1%** increase in private GDP if the other factors remain unchanged in the long run. This result is consistent with the theory of the relationship between FD and economic growth. This finding is also consistent with those of previous studies on Saudi Arabia (Altaee, Jafari, and Khalid 2016). An increase in broad money is associated with, among other things, the expansion of financial services and new financial products, both of which lead to satisfactory levels of financial products and liquidity in the real sector's activities.

Equation (5') shows that if BCP, another measure of FD, increases by **1%,** then private GDP increases by **0.1%** in the long run, ceteris paribus. The theory of financial intermediation considers BCP one of the key enablers for developing small and medium-sized enterprises. Even in the case of developing and emerging economies, bank loans can act as an essential source of capital for small and medium-sized companies. Our finding here shows that this channel of financial intermediation works for developing the private sector in the Saudi economy. This finding is also supported by the existing research on the nexus between FD and economic growth conducted in Saudi Arabia (Alshammary 2014; Ibrahim 2013; Masoud and Hardaker 2014; Nasir and Ali 2014; Osman 2014).

It appears that the magnitude of the elasticity of FD is relatively smaller than are those of labor and capital, as the magnitude of the elasticity of the FD is approximately **0.1,** regardless of whether the FD is measured by broad money size or BCP. These findings might imply that there is room for financial intermediation to play a more significant role in the development of the private sector. The Saudi policy strategy agenda also supports the further development of financial intermediation, as it is one of the key issues highlighted in Saudi Vision 2030. A standalone Vision realization program—the FSDP—aims to enable financial institutions to support private sector growth and outlines key targets that should be achieved by 2025.

According to the analysis, the data support the assumption that the long-run elasticities of private GDP concerning private labor and private capital are **0.52** and **0.48**, respectively. This result means that a 10% increase in the above two production factors is associated with a 5.2% and 4.8% increase in private sector growth, respectively. The theory of the production function is very well known in macroeconomics. Additionally, as discussed in Appendix 2B1, existing studies have estimated similar labor and capital elasticities of nonoil GDP for Saudi Arabia. Given these two points, we do not discuss the elasticities here. It is commonly accepted that service sectors are more labor intensive than are other sectors. In this regard, on average, during the period under consideration, 1970-2018, the shares of the service sector (excluding government services and utilities) and tradable sectors (agriculture and manufacturing) in private GDP were 66.7% and 23.9%, respectively. Moreover, the financial, insurance, and other business service sectors alone each had the largest share of private GDP, 32.1%, during the same period. This finding could explain why the degree of labor elasticity is higher than is that of capital elasticity according to our estimations.

In the analysis, we use the time trend to proxy for unknown and known variable factors (such as technological progress and total factor productivity) other than FD, labor, and capital, which are not explicitly included in the analysis. The estimations show that such unknown and known factors caused the growth of private GDP during the period under consideration. Finally, the estimated positive intercepts in Equations (4’) and (5’) imply that the technology of production goods and services, i.e., transforming inputs into outputs in the Saudi private sector, was progressive over the period.

The estimated SoA coefficients indicate that the short-run deviations of private GDP from the long-run relationship it establishes with its fundamentals (such as labor, capital, and FD) are temporary and will restore the long-run equilibrium relationship, thus implying that any policy or other shocks to the private sector would not result in a permanent change and that the sector’s development path will converge to its long-run equilibrium path. Numerically, almost **20%** of the remaining disequilibrium is corrected back to the equilibrium each year after a shock occurs if FD is measured by broad money size. This adjustment speed is increased eightfold if we consider BCP a measure of FD. These magnitudes of the SoA coefficients can be explained both economically and statistically. Economically, it is obvious that the impact of broad money on private sector development is different from that of BCP. Apparently, the adjustment process is considerably faster in the case of broad money, which is far more extensive than is BCP. Statistically, this difference in these factors can occur due to them having different specifications for private GDP and capital stock and different sets of restrictions imposed, which are nonlinear in nature, as reported in Tables 6 and 8.

As presented in Appendices 2B1 and 2C1, the other two identified long-run relationships are those for private labor and private capital stock. We do not discuss them in detail here because they are not the direct interests of the present study. Table 6 reports that a **1%** rise in private sector GDP leads to a **0.63%** increase in private employment if the variables in Equation (4) are considered to identify the long-run relationships. The same coefficient, i.e., the private GDP elasticity of private employment, is estimated to be **0.59** if the variables in Equation (5) are considered. The fact that the estimated magnitudes of the elasticities are quite similar may indicate their reasonableness. The positive output/income effects of employment are theoretically predicted because economic activities require an increasing number of workers as income grows. This result is also consistent with the findings of previous employment studies for Saudi Arabia (e.g., Hasanov et al. 2021a inter alia). Turning to the estimated long-run relationships for private capital stock, Table 6 reports that a **1%** increase in private GDP results in a **0.79%** expansion of private capital stock. Theoretically, this finding is consistent with the accelerator principle, which states that increased income/output can allow for increased investment and a desired level of capital. Previous research findings also support that output/income positively affects investments and capital stock (e.g., Bjerkholt 1993; Hasanov et al. 2020; Javid et al. 2021; Looney 1986). Table 8 reports the results for another representation of capital stock: a **1%** increase in BCP translates to a **0.34%** expansion of private capital stock in the long run. This finding also aligns with the accelerator principle, as the private sector will borrow from banks to achieve a desired level of capital. The time trend, representing unknown and known variable factors other than those included in the long-run equations, shows that such factors increase private labor and capital stock at the same magnitude, **1%** annually. Other noteworthy findings from the estimations are that the SoA coefficients (-0.46 and -0.27 in Table 6 and -0.22 and -1.03 in Table 8 for labor and capital, respectively) are significant and within a reasonable range, indicating that if policy and other shocks cause labor and capital to deviate from their long-run relationships, then these deviations will be temporary, and both variables will adjust to their long-run equilibrium relationships.

9. Concluding Remarks and Policy Insights

The FSDP is a vision realization programs, showing the great importance provided by Saudi Vision 2030, the masterplan of the economic development strategy of the Kingdom, to FD. The program was launched in 2017 to enable and support financial institutions to promote the development of the private sector. Against this backdrop, we investigate the role of FD in private economic growth in Saudi Arabia. We apply a multivariate cointegration method to the Saudi data in the extended production function framework for 1970-2018. We find that FD positively affects private sector development, with the elasticity being approximately 0.1 regardless of whether it is measured by a broad money-to-GDP ratio or BCP.

The insights from this study might be worth considering for authorities in their decision-making process. The key policy insight of this research is that decision-makers should consider the role of financial intermediaries in developing the private sector in conjunction with the developments of the labor and capital markets. In addition, decision-makers may consider that financial market components such as commercial BCP can boost the development of the private sector. Therefore, the loan process should be facilitated so that more private sector entities can obtain bank lending. We should not suggest the lowering of the lending rate, as the Saudi interest rate follows the US Federal Reserve rate due to the fixed exchange rate regime of the riyal since 1987. Therefore, administratively lowering interest or deposit rates would place depreciation pressure on the riyal-USD exchange rate. In contrast, Hasanov et al. (2022), inter alia, stated that the fixed exchange rate regime serves greatly for the development of the Saudi economy. To make the financial sector more supportive of the development of the private sector, the following policy measures should be considered, as highlighted in the FSDP (2021): enhancing the depth and breadth of the financial services and products offered, building an innovative financial infrastructure, managing risks through a thriving insurance sector, and enhancing the capabilities of the talent force.

Another policy insight that decision-makers can consider is that there is room for the development of the financial sector, as we find that the private sector effects of FD are relatively small in magnitude regardless of the measures considered. SV2030 also supports this insight, as the FSDP (2021) highlights a set of measures with which to develop the Saudi financial market. These measures include but are not limited to ensuring the formation of a developed capital market and promoting and enabling financial planning, alongside its numerical targets such as increasing the share of financing for small and medium-sized enterprises from banks to 20% in 2030, compared to 5.7% in 2019; increasing the market value of the stock market as a percentage of GDP from 66.5% in 2019 to 88% in 2030 (excluding the Aramco IPO); and achieving total banking assets of 4.553 billion riyals in 2030 from 2.631 billion riyals in 2019.

Endnotes

[1] Here and hereafter, private sector, private GDP, and labor and capital refer to those of the nonoil sector.

[2] There are mixed approaches to calculating the broad money supply share of GDP in terms of considering nominal or real values. For example, Levine (1997) and Gelb (1989) considered nominal money supply and real GDP, Calderon and Liu (2003) considered real money supply and real GDP, and Arestis and Demetriades (1997) considered nominal money supply and nominal GDP. We use nominal money supply and real GDP here following seminal studies such as Levine (1997) and Gelb (1989), among others. Note, however, that we find that the empirical (estimations and tests) results are very similar to each other, regardless of whether we consider nominal money supply or real money supply share in real GDP.

[3] FD is proxied by the ratio of the liquid liabilities of the financial system to GDP, which for most countries equals M2/GDP. King and Levine (1993) showed that this measure is closely associated with long-run growth.

[4] The problem of omitted variable bias can arise if the equilibrium correction term(s) of the other cointegrating relationship(s) enters the short-run specification of the variable of interest in a significant way (see, e.g., Badinger 2004; Dibooglu and Enders 1995; Enders 2015; Ericsson and MacKinnon 2002).

[5] We include both intercept and trend in the test equation and select the maximum lag length to be 3, as we do for the other variables in the standard ADF and PP tests. We consider 1975 as the break date based on the graphical illustrations of k and Δk in Figure 3. Finally, we consider that the break is sudden, i.e., an additive outlier.

[6] The estimation sample covers 1975-2018, as our data start in 1970; we set the maximum lag order to four in the VAR model and consider the first difference of the variables in the VECM.

[7] Note that reducing the lag length from three to two causes severe serial correlation in the residuals of the VAR equations.

[8] Sometimes, it is stated in the literature that SoA should not be smaller than negative one. However, we believe that it should be in the range of (-2,0). As explained in Enders (2015, 374, 377–378), the concept of SoA is the same as the autoregressive coefficient in the ADF unit root test equation being in first-differenced form. Enders (2015, 205) shows that stationarity holds if the autoregressive coefficient is in the interval of (-2,0). Note that other empirical studies also have found SoA coefficients to be smaller than negative one (e.g., see Juselius 2006, 249; Loayza and Ranciere 2005; Narayan and Smyth 2006; Olczyk and Kordalska 2017; Shittu, Yemitan, and Yaya 2012); for Saudi Arabia, see Hasanov (2021), Hasanov et al. (2022), and Hasanov et al. (2021a).

[9] We do not need to include any of the previous dummy variables in the VAR model here, as its residuals do not demonstrate any significant outliers and are well-behaved. Additionally, including dummies causes serial correlation in the residuals.

[10] Panel B indicates a kurtosis issue from the joint normality test. First, as discussed in the literature, kurtosis is not a serious issue compared to the skewness issue (see the discussions in Hendry and Juselius 2001, inter alia). Second, this issue stems mainly from the residuals of the l equation. The residuals of the y equation, our main interest, show neither kurtosis nor skewness, as the sample χ^2 values (and their probabilities) for the null hypotheses of no skewness and no kurtosis are 0.89 (0.35) and 2.83 (0.09), respectively.

[11] We estimate the VAR model of y, $m2$, k, and l with three lags and the VAR model of y, k, l, and bc with four lags. We include intercept and trend as well as three pulse dummy variables (taking unity in 1983, 1986, and 2002, respectively, and zero otherwise) in the former VAR model, while the latter VAR model includes only intercept and trend to obtain the same estimation results as those reported in Tables 5 and 7, respectively. Additionally, following the methodology in Doornik and Hendry (2018), Section 8.9, we clear the unrestricted status of intercept and trend in specifying the VAR models. The estimation period is 1975-2018.

[12] To impose no rank restrictions, we set the rank of cointegration to four (i.e., equal to the number of economic variables) following the methodology used in Doornik and Hendry (2018), Section 4.7. The estimation period is 1975-2018.

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Appendix 1. Econometric Methods

In this section, we describe the econometric methods used for the unit root and cointegration tests and for estimating the long-run parameters. This section briefly introduces these unit root (UR) tests and then describes the Johansen cointegration method.

Unit Root Tests

The majority of economic indicators trend over time stochastically. Hence, it is essential first to check their stationarity through UR tests to prevent spurious results. The most widely employed UR tests are the augmented Dickey-Fuller (ADF) test (Dickey and Fuller 1979) and the Philips-Perron (PP) test (Phillips and Perron 1988), although there are many UR tests available.

The ADF equation for a given variable, y_t, can be written as follows in the case of an intercept and trend:

$$\Delta y_t = b_0 + b_1 t + b_2 y_{t-1} + \sum_{i=1}^{l} \gamma_i \Delta y_{t-i} + v_t \tag{6}$$

where b_0 and t are a constant term and a linear time trend, respectively; l and Δ denote the number of lags and the first difference operator, respectively; and v_t refers to white noise errors. The ADF sample value is represented by the *t*-statistic on b_2. The null hypothesis of the UR is rejected if this value is smaller than the critical ADF values, in absolute terms, at different significance levels, meaning that y_t has a UR and therefore is not stationary. If this value is greater than the critical ADF values, in absolute terms, at different significance levels, then the null hypothesis can be rejected, which means that y_t is not nonstationary.

The only difference between the PP and ADF tests is that to remove the serial correlation problem in the residuals, the former uses nonparametric statistical methods but not lags of the dependent variable. A detailed discussion of the PP test can be found in Phillips and Perron (1988).

Johansen Cointegration Method

Cointegration theory articulates that if variables are nonstationary and their integration orders are the same, usually one, then it is meaningful to check whether they have a long-run relationship using cointegration test(s). Again, cointegration theory states that if n number of variables are under consideration, then there can be *n-1* number of cointegrating relationships at maximum. However, only the Johansen cointegration test can discover the number of cointegrating relationships among the variables if there is more than one variable (Enders 2010; Engle and Granger 1987; Johansen 1988). Therefore, we prioritize the Johansen method, as we have more than two variables in our analysis.

The full information maximum likelihood of the vector error correction model (VECM) of Johansen (1988) and Johansen (1991) can be expressed as follows:

$$\Delta z_t = \Pi z_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \, \Delta z_{t-i} + c + e_t \tag{7}$$

where z_t is a (n × 1) vector of the n endogenous/modeled variables, c is a (n × 1) vector of constants, Γ represents a (n × (k – 1)) matrix of short-run coefficients, e_t denotes a (n × 1) vector of white noise residuals, and Π is a (n × n) coefficient matrix. If matrix Π has reduced rank (0 < r < n), then it can be split into a (n × r) matrix of loading coefficients α and a (n × r) matrix of cointegrating vectors β. The former indicates the importance of the cointegration relationships in the individual equations of the system and of the speed of adjustment to equilibrium, while the latter represents the long-term equilibrium relationship; thus, $\Pi = \alpha\beta'$.

Johansen's reduced rank regression approach of testing for cointegration estimates matrix Π in an unrestricted form first and then tests whether the restriction implied by the reduced rank of Π can be rejected. In particular, the number of independent cointegrating vectors depends on the rank of Π, which in turn is determined by the number of its characteristic roots that are different from zero. Max-eigenvalue and trace test statistics are used to test for nonzero characteristic roots.

Note that significance, multivariate stationarity, and weak exogeneity tests are usually conducted in the Johansen framework using the estimated VECM (Johansen 1992a, 1992b). If a given variable, X, in the long-run space is significant, then the null hypothesis expressing that its long-run coefficient β_X is zero can be rejected at conventional levels. The multivariate stationarity or trend stationarity of a given X variable can be expressed with the null hypothesis that its long-run coefficient β_X is unity, while the long-run coefficients of other explanatory variables are zero. If the null hypothesis cannot be rejected, then variable X is considered (trend) stationary. The rejection of the null hypothesis indicates that X is a unit root process. The weak exogeneity of a given variable, X, implies that the null hypothesis of α_X being zero cannot be rejected. The weak exogeneity indicates that the disequilibrium of the long-run relationship does not feed back to the given variable X's equation. If the null hypothesis of α_X being zero can be rejected, then X is not a weakly exogeneous variable, meaning that the disequilibrium of the long-run relationship feeds back to its equation.

Small Sample Bias Correction in the Johansen Method

Johansen (2002) discusses that in the case of small samples, the max-eigenvalue or trace test statistics can be biased to reject the null hypothesis of no cointegration. Regarding this issue, Reinsel and Ahn (1992) suggest $\frac{T-kn}{T}$ correction to the trace and max-eigenvalue test statistics, where k is the lag length of the underlying vector autoregressive (VAR) model in levels and n and T are the number of endogenous variables and observations, respectively.

Appendix 2. Results of Empirical Analysis

2.A. Unit Root Test

Table 4 reports that the null hypothesis of the unit root cannot be rejected for the FD measures of *bc* and *m2*, given that the sample values are smaller than the respective critical values in absolute terms. The null hypothesis can be rejected at the 1% level for the first differences of these variables, i.e., *Δbc* and *Δm2*. This result drives us to conclude that these variables are an integrated order of one type of process, that is, I(1) processes. This conclusion holds regardless of whether the ADF or PP test results are considered.

Although the ADF test results suggest that *y* is a trend-stationary process, the PP test results indicate that it is a unit root process. The time path of the variable in Panel A of Figure 3 illustrates that it is most likely a unit process rather than a trend-stationary process, as it is difficult to see a deterministic trend in the development of the variable, around which it has a constant mean and/or variance. To this end, it can be concluded that *y* is a unit root process. Both test results agree that *Δy* is a stationary process. Hence, the variable can be considered an I(1) process. The PP test results suggest that *l* and *k* are unit root processes, although the ADF results indicate a trend-stationary process with very weak evidence (only at the 10% significance level, but at the 1% and 5% significance levels, the ADF results also suggest a unit root process).

Additionally, the time trajectories of the variables do not suggest that they are trend-stationary processes since they follow a similar pattern as that of *y* (see Panel A of Figure 3). Both test results strongly reject the null hypothesis for *Δl*; thus, we can conclude that *l* is also an I(1) process. For *Δk*, both test results reject the unit root process only at the 10% significance level. This leads us to mixed conclusions: *k* may be stationary at the second difference at the higher significance level, meaning that it is an integrated order of two, that is, the I(2) process, which one may expect for capital stock. Alternatively, *k* may be an I(1) process, meaning that *Δk* is stationary with a structural break, as can be predicted from its time trajectory in Panel B of Figure 3. To this end, we run the ADF test with a structural break on *Δk* to make a robust decision concerning the integration order of *k*. Since the ADF test results weakly suggest that *k* might be a stationary trend process, we also perform the ADF test with a structural break on *k*. For *k* and *Δk*, the test values are -1.737 and -4.413, respectively. Comparing these sample values with the critical values of -4.4, -3.8, and -3.5 at 1%, 5%, and 10%, respectively, one can conclude that *k* is a unit root process and that *Δk* is a stationary process, leading to the conclusion that *k* is an I(1) process.[5] In summary, our conclusion from the unit root tests is that all the variables follow the I(1) process.

2.B. Specifying a VAR Model/VECM and Cointegration Analysis for Equation (4)

Following Johansen's method (see Johansen 1988; Johansen 1991; Juselius 2006), we first specify a VAR of the four endogenous variables (*y, m2, l, and k*) with a lag order of four as a maximum.[6] We include intercept and trend in the VAR model as exogenous variables. We also include three dummy variables, taking unity in 1983, 1986 and 2002 and zero for other years based on the inspection of the residuals of the VAR model. The first two dummy variables capture an enormous decline in the residuals of the y and I equations, mostly caused by the economic recession that happened in 1981-1986. The last dummy variable captures a large jump in 2002 in the residuals of the m2 equation, as M2 grew by more than 14%, while GDP declined by 3% due to an oil production cut of 10% that year. Then, we perform the lag exclusion test and use lag order selection criteria to identify the optimal lag order. The lag exclusion test indicates that four lags can be reduced to three without losing information for the y, I, and k equations. This approach can also be taken for the *m2* equations at the 5% significance level. Hence, the joint significance of four lags can be reduced to three lags for all four equations without information loss and can be accepted at the 5% significance level as the sample. The χ^2 value and associated p value are 27.60 and 0.04, respectively. The hypothesis that three lags can be reduced to two lags without any information loss can be profoundly rejected, as the sample χ^2 value and associated p value are 44.94 and 0.00, respectively. Regarding the lag order selection criteria test, the likelihood ratio and Schwarz criterion prefer three lags, while final prediction error, the Akaike information criterion, and Hanna-Quinn criterion indicate four lags as an optimal lag order. We decide to select three lags as the optimal order. The residuals of the VAR with three lags do not have any problems in terms of serial correlation, nonnormal distribution, or heteroscedasticity and are also stable over time, as documented in Table 5, Panels A through D.[7]

Table 5. VAR residual diagnostics and cointegration test results for Equation (4).

Panel A: Serial Correlation LM Test[a]

Lags	LM-Statistic	P value
1	24.45	0.08
2	12.95	0.68
3	18.19	0.31

Panel E: Johansen Cointegration Test Summary

Data Trend:	None	None	Linear	Linear	Quadratic
Test Type:	(a) No ***C*** and ***t***	(b) Only ***C***	(c) Only ***C***	(d) ***C*** and ***t***	(e) ***C*** and ***t***
Trace:	3	3	2	3	3
Max-Eig:	3	3	2	3	3

Panel B: Normality Test[b]

Statistic	χ^2	d.f.	P value
Skewness	4.63	4	0.33
Kurtosis	3.47	4	0.48
Jarque-Bera	8.10	8	0.42

Panel F: Johansen Cointegration Test Results for Type (d)

Null hypothesis:	$r = 0$	$r \leq 1$	$r \leq 2$	$r \leq 3$
λ_{trace}	105.25***	65.03***	33.70***	8.09
λ^a_{trace}	76.54***	47.29**	24.51*	5.88
λ_{max}	40.22***	31.33***	25.61***	8.09***
λ^a_{max}	29.25	22.78	18.63	5.88

Panel C: Heteroscedasticity Test[c]

White	χ^2	d.f.	P value
Statistic	267.54	290	0.82

Panel D: VAR Stability Test

Root Modulus
0.91-0.06i0.91
0.91 + 0.06i0.91
0.75-0.41i0.85
0.75 + 0.41i0.85

Notes: [a] The null hypothesis in the serial correlation LM test is that there is no serial correlation at lag order h of the residuals; [b] The system normality tests with the null hypothesis of the residuals are multivariate normal;[c] The White heteroscedasticity test takes the null hypothesis of there being no cross-term heteroscedasticity in the residuals; χ^2denotes chi-squared; d.f. denotes degrees of freedom; and ***C*** and ***t*** indicate intercept and trend. r is the rank of matrix Π, i.e., number of cointegrated equations; λ_{trace} and λ_{max} are the trace and max-eigenvalue statistics, while λ^a_{trace} and λ^a_{max} are the adjusted versions of them, respectively; *** , **, and * denote the rejection of the null hypothesis at the 1%, 5% and 10% levels; critical values for the cointegration test are taken from MacKinnon, Alfred, and Leo (1999); and the estimation period is 1975-2018.

Therefore, we opt for the VAR with three lags for further tests and estimations. We transform the VAR model into a VECM with two lags to conduct the Johansen cointegration test.

2.B.1. Imposed Restrictions and Identification of Long-Run Equations

Although we report the cointegration test results for all five possible versions in Panel E of Table 5, social and economic processes are usually better represented by versions (c) and (d). We prefer the latter to the former, as our theoretical model contains a time trend. Panel F of Table 5 reports standard and adjusted trace and max-eigenvalue sample statistics for version (d). In their unadjusted form, both test statistics reject the null hypothesis of three cointegrating equations against two at the 1% significance level. Even the adjusted trace statistics reject the null hypothesis of three cointegrating equations against two. The adjusted max-eigenvalue sample statistics cannot reject the null hypothesis of no cointegrating equation, but this finding is difficult to believe because we have a theoretical reason for there to be at least one long-run relation among the variables. In conclusion, we conclude that there are three cointegrating equations.

Statistically, the existing three cointegrating vectors require imposing at least three restrictions on each of the equations to identify the long-run equations (e.g., Pesaran and Shin 2002). The Johansen method, by default, imposes one unity and two zero restrictions in a diagonal manner. However, this way of imposing restrictions is not helpful for us, as it assumes that for example, in the first equation, Equation (4), the coefficients of *m2* and *l* are zero. Even if we change the orders of the explanatory variables in Equation (4) and thereby in the VAR model/VECM, then imposing, by default, restrictions will put zero restrictions on one of the economic variables, as it needs to put one unity and two zero constraints on five variables, including the time trend. This identification approach has been criticized as ‘pure mathematical convenience’ by Pesaran and Shin (2002), who, instead, suggested the use of a theory-guided method to identify the long-run equations. This theory-guided approach takes Johansen’s just identified cointegration vectors as given and replaces ‘statistical’ restrictions with economically meaningful restrictions. Then, additional theoretically meaningful restrictions can be imposed on the just identified equations, and the χ^2 test is used to check whether overidentifying restriction(s) is (are) valid (see Pesaran and Shin 2002; Zou et al. 2004 inter alia). We follow this approach in imposing restrictions on the cointegrating vectors to identify our long-run equations in theoretically meaningful ways. As discussed in the literature, identifying the long-run equations in an economically meaningful way is not easy when the number of cointegrating relations is more than one. Such identification also requires a great deal of time to validate the multiple options dictated by economic theory, data evidence, and the country’s stylized facts. This challenge is further aggravated by three additional statistical constraints: (i) the imposed theoretical framework should yield significant long-run coefficients, (ii) numerous restrictions (which total three in our case) for just identification on each cointegrating vector has to be respected and (iii) all restrictions on long-run and loading coefficients from all VECM equations have to be significantly binding. Table 6 reports the imposed restrictions on the long-run and loading (speed of adjustment) coefficients and their statistical validity, as selected among many options.

Appendix 2. Results of Empirical Analysis

Table 6. Long-run estimation, identification, and test results for Equation (4).

Panel A: Identified long-run equations for *y*, *l*, and *k* and their SoA coefficients					
Long-run equation					SoA coefficient
$\widehat{y}_t = 1.96 +$	$0.10\ m2_t +$	$0.52\ l_t +$	$0.48\ k_t +$	$0.003\ t_t$	$\alpha_{y/y} = -1.60\ (0.29)$
	(0.02)			(0.001)	
$\widehat{l}_t = 0.22 +$	$0.63\ y_t +$	$0.01\ t_t$			$\alpha_{l/l} = -0.46\ (0.09)$
	(0.04)	(0.002)			
$\widehat{k}_t = 3.30 +$	$0.79\ y_t +$	$0.01\ t_t$			$\alpha_{k/k} = -0.27\ (0.04)$
	(0.04)	(0.002)			
Panel B: Imposed restriction on long-run and SoA coefficients [a]					
$\beta_{y/y} = 1$	$\beta_{y/l} = 0.52$	$\beta_{y/k} = 0.48$			
$\beta_{l/l} = 1$	$\beta_{l/m2} = 0$	$\beta_{l/m2} = 0$			
$\beta_{k/k} = 1$	$\beta_{k/m2} = 0$	$\beta_{k/m2} = 0$			
$\alpha_{y/l} = 0$	$\alpha_{y/k} = 0$				
$\alpha_{m2/l} = 0$					
χ^2 [p value]: **0.27 [0.97]**					

Notes: [a] The null hypothesis is that a given restriction is nonsignificant; values in parentheses and brackets are standard errors and probability values, respectively. $\widehat{Y}_t$ means estimated/fitted Y_t. $\beta_{X/Y}$ indicates the long-run coefficient of *Y* in the long-run *X* equation. $\alpha_{X/Y}$ indicates the loading coefficient of the disequilibrium of the long-run Y equation in the equilibrium *X* correction equation. The estimation period is 1975-2018.

For the first equation, which is for *y* and is our main interest, we try to keep all the explanatory variables in Equation (4) if such a framework is statistically supported. Recall that the by-default approach to identification does not allow us to achieve this aim. Hence, we use a theory-guided approach to identification following Pesaran and Shin (2002) and Zou et al. (2004). To this end, we first impose various elasticities of *y* with respect to *l* and *k* considering recent studies conducted for Saudi GDP or nonoil GDP, such as Aljebrin (2013), Hasanov et al. (2019), and Hasanov et al. (2021b), as well as the references therein. The imposed values of 0.52 and 0.48 for labor and capital elasticities, respectively, cannot be rejected statistically according to the sample value of χ^2. Additionally, these restrictions provide (i) identification of the long-run *y* equation with economically interpretable and significant elasticity estimates for the FD measures, (ii) meaningful and significant estimates for the SoA coefficient in the short-run *y* equation, and (iii) support for the identification of the other two equations for *l* and *k* with significant and economically interpretable estimates of the long-run and SoA coefficients. Importantly, just identification is not present for the first equation, leaving the other two equations not identified because doing so would lead to the whole cointegration space not being identified. Hence, we normalize the second equation for *l* and the third equation for *k* and impose restrictions after checking the statistical validity of other normalizations and restrictions. For the *l* equation, we check different assumptions considering *l* to be dependent on different combinations of *y*, *k*, *m2* and the *time trend* and test the significance of each of these assumptions. The best option among them is the option that assumes *l* to be a function of *y,* a measure of economic activity and *time trend* representing the impact of other factors. This is a reasonable assumption given that economic activity and wage rate are theoretically the main drivers of employment. For *k*, the best option both theoretically and statistically is the same as that for the *l* equation, that is, considering *y* and the *time trend* as long-run explanatory variables. Table 6, Panel B, presents restrictions that we impose on the long-run and loading (i.e., SoA) coefficients to identify these equations. The sample value of the χ^2 test is 0.27, with a probability value of 0.97, meaning that the imposed restrictions cannot be rejected and that they are held with high probability. Panel A reports the identified long-run equations for *y*, *l*, and *k* and their respected SoA coefficients, which are shown to be significant, as the panel reports. Additionally, the SoA coefficients are negative and in the reasonable range of (-2; 0).[8] This finding indicates that although *y*, *l*, and *k* deviate from their established long-run relationships in the short run, these deviations are temporary, and the variables can return to their long-run relationships. Imposing (zero) restrictions on the SoA coefficients can be understood as testing for the weak exogeneity of a given variable, that is, to examine whether long-run disequilibrium terms with a one-year lag enter the short-run (equilibrium correction) equation of a given variable in a significant way. We check the weak exogeneity of other variables in the estimated VECM, but they do not produce reasonable results regarding the theoretical and statistical aspects of the estimations. Finally, the estimated VECM with the imposed restrictions reported in Table 6 successfully passes postestimation tests, including serial correlation, normality, heteroscedasticity, and cointegration tests of version (d), still indicating the presence of three cointegrating equations.

2.C. Specifying a VAR Model/VECM and Cointegration Analysis for Equation (5)

Following the same methodological procedures as those for Equation (4) above, we set up a VAR model of the endogenous variables of *y, bc, l,* and *k* with a maximum lag order of four while considering intercept and trend exogenous variables.[9] A joint significance lag exclusion test indicates that four lags cannot be reduced to 3 lags without loss of information, as the sample χ^2 value and associated p value are 32.60 and 0.01, respectively. In addition, the final prediction error, Hanna-Quinn criterion, and Akaike information criteria prefer four lags, while the likelihood ratio uses three lags, and the Schwarz criterion uses two lags. Thus, both tests favor mainly the lag length of 4 as an optimal lag order. As an additional check, we estimate the VAR model with three lags, which leads to the skewness problem in the joint test of normality of all equations' residuals and the individual test of normality for the residuals of the *y* equation, our main interest. The same skewness problem occurs for the case of the VAR model estimated with two lags. Thus, from the standpoint of well-behaved residuals, one should select four lags, although doing so would consume 16 degrees of freedom (without considering the deterministic regressors). The residuals of the VAR model with four lags do not have any serial correlation, nonnormal distribution, or heteroscedasticity problems. This model is also stable over time, as documented in Panels A through D of Table 6.[10]

Table 7. VAR model residual diagnostics and cointegration test results for Equation (5).

Panel A: Serial Correlation LM Test[a]		
Lags	LM-Statistic	P value
1	24.45	0.47
2	12.95	0.89
3	18.20	0.14

Panel E: Johansen Cointegration Test Summary					
Data Trend:	None	None	Linear	Linear	Quadratic
Test Type:	(a) No C and t	(b) Only C	(c) Only C	(d) C and t	(e) C and t
Trace:	2	3	2	3	3
Max-Eig:	2	3	2	3	3

Panel B: Normality Test b			
Statistic	χ^2	d.f.	P value
Skewness	4.17	4	0.38
Kurtosis	15.45	4	0.00
Jarque-Bera	19.62	8	0.01

Panel F: Johansen Cointegration Test Results for Type (d)				
Null hypothesis:	$r = 0$	$r \leq 1$	$r \leq 2$	$r \leq 3$
λ_{trace}	105.31***	55.97***	26.62**	5.05
λ^a_{trace}	57.44	30.53	14.52	2.75
λ_{max}	49.34***	29.35**	21.57**	5.05
λ^a_{max}	26.91	16.01	11.77	2.75

Panel C: Heteroscedasticity Test [c]			
White	χ^2	d.f.	P value
Statistic	332.99	340	0.60

Panel D: VAR Stability Test
Root Modulus
0.91-0.21i0.93
0.91 + 0.21i0.93
0.73-0.51i0.89
0.73 + 0.51i0.89

Notes: [a] The null hypothesis in the serial correlation LM test is that there is no serial correlation at lag order h of the residuals; [b] The system normality test with the null hypothesis of the residuals being multivariate normal; [c] The White heteroscedasticity test takes the null hypothesis of no cross terms heteroscedasticity in the residuals; χ^2 denotes chi-squared; d.f. means degree of freedom; and *C* and *t* indicate intercept and trend. *r* is the rank of matrix Π, i.e., number of cointegrated equations; λ_{trace} and λ_{max} are the trace and max-eigenvalue statistics, while λ^a_{trace} and λ^a_{max} are their adjusted versions, respectively; *** , **, and * denote the rejection of the null hypothesis at the 1%, 5% and 10% levels; critical values for the cointegration test are taken from MacKinnon, Alfred, and Leo (1999); and the estimation period is 1975-2018.

Therefore, we specify the VAR model with four lags and transfer it into a VECM with three lags for our further tests and estimations.

2.C.1. Imposed Restrictions and Identification of Long-Run Equations

Three out of five possible options suggest three cointegration equations, as Panel E of Table 6 presents. In particular, in version (d), which includes the time trend, the trace and max-eigenvalue statistics indicate three cointegrating equations. Panel F of Table 6 reports standard and adjusted trace and max-eigenvalue sample statistics for version (d). In their unadjusted form, both test statistics reject the null hypothesis of three cointegrating equations against two. The null hypothesis of no cointegrating equation cannot be rejected by either of the adjusted trace and max-eigenvalue sample statistics. However, it is not straightforward to accept the suggestion that there is no cointegration among the variables in Equation (5), as one would theoretically expect a long-run relationship, as has been found in the empirical studies in the Literature Review section. Moreover, we conclude above that three theoretically interpretable and statistically acceptable long-run relationships exist among the variables in Equation (4). Thus, as a research decision, we also conclude that three cointegrating equations exist among the variables in Equation (5).

We need to impose at least three restrictions on each of these three equations to obtain just identified cointegration space. We place many different restrictions on the cointegrating equations using the abovementioned theory-guided approach. Table 8 reports the final set of imposed restrictions on the long-run and SoA coefficients as well as their statistical validity. As reported, for the final set of restrictions, we obtain an $\chi 2$ value of 3.1, with a p value of 0.4, indicating that the imposed restrictions cannot be rejected.

We try to keep all the explanatory variables in the first equation, as it is our main equation of interest. This attempt works theoretically and statistically for the explanatory variables of *bc*, *l,* and *k* but not for that of *time trend*. The coefficient of *time trend* appears nonsignificant and, additionally, takes a negative sign, which is difficult to explain, as the overall development of private GDP did not decline over the period considered. Thus, we place a zero/exclusion restriction on *time trend* in *the y equation, the cointegration space's first equation*. The following is our explanation for why the time trend works in Equation (4), as reported in Table 6, but not Equation (5), as reported in Table 8. In the former case, we use *m2*, which is a derivative variable

Table 8. Long-run estimation, identification, and test results for Equation (5).

Panel A: Identified long-run equations for *y*, *l*, and *k* and their SoA coefficients

Long-run equation				SoA coefficient
$\widehat{y}_t = 0.69\ +$	$0.10\ bc_t +$	$0.52\ l_t\ +$	$0.48\ k_t$	$\alpha_{y/y} = -0.20\ (0.08)$
	(0.01)			
$\hat{l}_t = 0.69\ +$	$0.59\ y_t\ +$	$0.01\ t_t$		$\alpha_{l/l} = -0.22\ (0.10)$
	(0.04)	(0.002)		
$\widehat{k}_t = 9.04\ +$	$0.34\ bc_t\ +$	$0.01\ t_t$		$\alpha_{k/k} = -1.03\ (0.17)$
	(0.05)	(0.005)		
Panel B: Imposed restriction on long-run and SoA coefficients [a]				
$\beta_{y/y} = 1$	$\beta_{y/l} = 0.52$	$\beta_{y/k} = 0.48$	$\beta_{y/t} = 0$	
$\beta_{l/l} = 1$	$\beta_{l/bc} = 0$	$\beta_{l/k} = 0$		
$\beta_{k/k} = 1$.	$\beta_{k/y} = 0$.	$\beta_{k/l} = 0$		
$\alpha_{y/l} = 0$				
$\alpha_{k/l} = 0$				
χ^2 [p value]: **3.13 [0.37]**				

Notes: [a] The null hypothesis is that a given restriction is nonsignificant. Values in parentheses and brackets are standard errors and probability values, respectively. $\widehat{Y}_t$ denotes estimated/fitted Y_t. $\beta_{X/Y}$ indicates the coefficient of Y in the long-run X equation. $\alpha_{X/Y}$ indicates the loading coefficient of the disequilibrium of the long-run Y equation in the equilibrium X correction equation. The estimation period is 1975-2018.

(the ratio of M2 to GDP) and has an uneven trend component (if we were illusionary, then the variable would be decomposed into a trend, a cycle, and irregular components) compared to that of *bc* used in the latter case (see Panel A of Figure 3). This uneven trend component of *m2* does not better capture the trend component of *y,* and hence, the time trend variable captured whatever aspects are excluded from *l*, *k* and *m2* in Table 6. However, the trend component of *bc* is not uneven and may better capture the trend component of *y*, which includes whatever aspect are excluded from *l* and *k*; hence, there is no need for an additional time trend variable to capture the trend component of *y*. We consider whether the elasticities of *y* with respect to *l* and *k* can be 0.52 and 0.48, respectively, as we do in Table 6 above. These restrictions are theoretically reasonable, as discussed above, and are not rejected, as Table 8 reports. Another reason for us to impose the same restrictions here as we do in Table 6 for Equation (4) is that if *y* is significantly dependent on *l* and *k* with the mentioned elasticities in Table 6, then the same should conceptually hold true for Equation (5), which would indicate that the estimates of the impacts of *l* and *k* on *y* are consistent and do not change across specifications. We estimate the elasticity of the FD measure for Equation (5) in Table 8 to be almost the same as that obtained for Equation (4) in Table 6. This finding may indicate that regardless of whether FD is measured by *m2* or *bc*, its numerical impact on private GDP is approximately 0.10.

For the second equation, we check whether *l* can be a function of *k* in a significant way, which theoretically assumes either complementarity or substitution between them to be determined by the sign of the coefficient of *k*. We also check that *l* is dependent on *bc*, which theoretically is in line with the FD-led economic development hypothesis. However, none of the assumptions are significant in the case of Saudi data. Hence, we end up with the same specification as that in Table 6, where *l* is a function of *y* and *time trend*. This specification is theoretically grounded given that the demand for labor is dependent mainly on output and wage rate. In addition, this theoretical framework for labor demand is supported by previous empirical research on the Saudi economy (see Hasanov et al. 2021a, inter alia). The specification has significant coefficients for *y* and *time trend,* as Table 8 presents.

Finally, in the search for a relevant representation of *k* using the variables in Equation (5), we check whether *bc* can be one of the significant explanatory variables given that conceptually, *bc,* as an investment in the private sector, leads to an increase in *k*. This assumption holds with the presence of a time trend and the absence of *y* and *l* in the *k* equation. This specification here is different from that in Table 6 for Equation (4). Therefore, this specification provides additional information explaining the behavior of *k* over the period considered. The magnitudes of the coefficients on the time trend variable in the *l* and *k* equations in Tables 8 and 6 are almost the same. This finding may indicate a consistent impact of the other variables, which are not included explicitly in Equations (4) and (5) and, thus, in the analysis, on *l* and *k* being approximately 0.01.

Additionally, Table 8 documents that the SoA coefficients on the disequilibrium terms from the identified long-run specifications of *y*, *l*, and *k* are significant with the expected negative signs. This means that the identified long-run relationships are stable, as shocks to them are temporary, and *y*, *l*, and *k* will be restored to the identified relationships. This result further indicates that as expected, the mentioned variables are not weakly exogenous to the disequilibrium terms of their identified long-run relationship. We also impose zero restrictions on the other SoA coefficients in the VECM. Among these restrictions, only the zero restrictions on $\alpha_{y/l}$ and $\alpha_{k/l}$ appear significant. In other words, it is found that the one-year lagged disequilibrium terms of the long-run *y* and *l* equations can be excluded from the short-run (equilibrium correction) *l* and *k* equations, respectively. This finding means that *l* and *k* are weakly exogenous to the long-run relationships of *y* and *l*, respectively, assuming that zero/exclusion restrictions on more than these two SoA coefficients causes problems, such as serial correlation, heteroskedasticity, and nonnormality in the residuals of the VECM. Otherwise, the residuals of the VECM are well behaved, and the cointegration test still indicates three cointegrating equations in version (d).

2.D. Parameter Constancy Tests for the Estimated VAR Model/VECM

In this section, we perform Chow tests, forecast tests, and recursive eigenvalue tests to check the estimated parameters' constancy in the VAR models using the PcGive package in OxMetrics 8 (see Doornik and Hendry 2018, Sections 8.9 and 4.7 for more information). We run the mentioned tests on the estimated VAR models for Equations (4) and (5), as reported in Tables 5 and 7, respectively.[11]

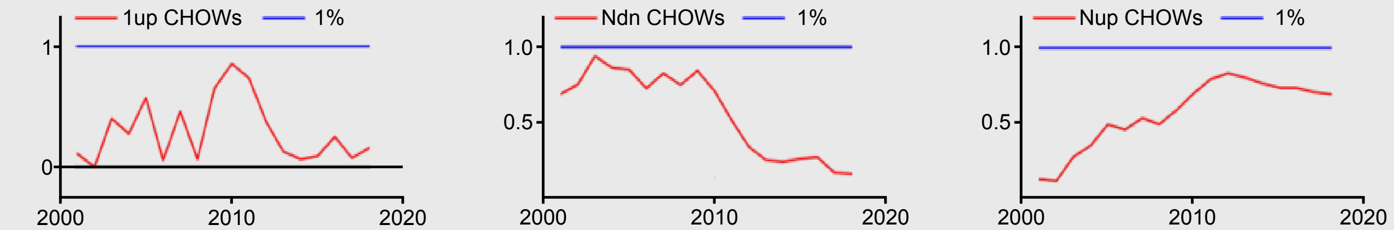
Figure 4 illustrates the results of the system-based 1-step Chow test, breakpoint Chow test, and forecast Chow test for the VAR models.

Apparently, none of the red lines cross the blue lines, indicating that *the null hypothesis of the parameters estimated in different samples being equal* cannot be rejected, meaning that the parameters of both VAR models are stable over time.

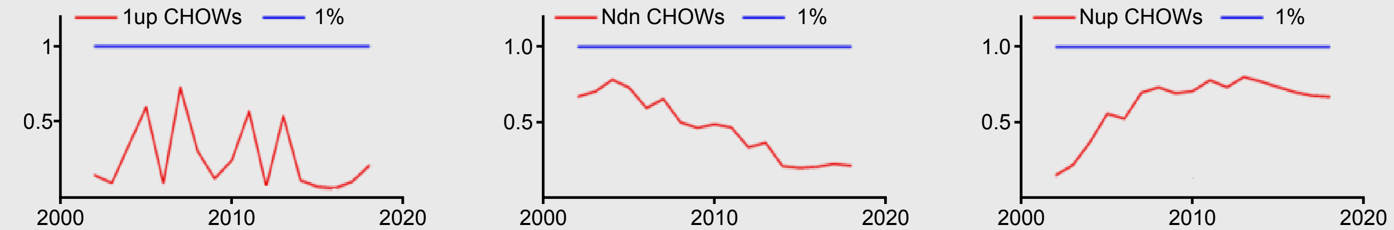
Next, we perform a forecast test for parameter constancy. To do so, we estimate both VAR models until 2015 and leave 2016-2018 for the forecasting horizon. Notably, this is a quite difficult exercise for these VAR models, as

Figure 4. Chow test results for parameter constancy.

Graph 4.A. Test results for the VAR estimated for equation (4)



Graph 4.B. Test results for the VAR estimated for equation (5)



Note: 1up CHOW = system-based 1-step Chow test, Ndn CHOWs = system-based breakpoint Chow test, Nup CHOWs = system-based forecast Chow test, and the blue line indicates significance at 1%.

Appendix 2. Results of Empirical Analysis

Table 9. Parameter constancy forecast tests.

Panel A. Test results for the VAR estimated for Equation (4)	
using Omega	Chi^2 (12) = 20.99 [0.05] F(12,24) = 1.75 [0.12]
using V[e]	Chi^2 (12) = 12.71 [0.39] F(12,24) = 1.06 [0.43]
using V[E]	Chi^2 (12) = 9.52 [0.66] F(12,24) = 0.79 [0.65]
Panel B. Test results for the VAR estimated for Equation (5)	
using Omega	Chi^2 (12) = 16.35 [0.18] F(12,23) = 1.36 [0.25]
using V[e]	Chi^2 (12) = 9.29 [0.68] F(12,23) = 0.77 [0.67]
using V[E]	Chi^2 (12) = 7.31 [0.84] F(12,23) = 0.61 [0.81]

domestic energy prices and fiscal reforms were implemented while international oil prices collapsed significantly in 2016-2018. In other words, models estimated until 2015 may not capture the changes that happened in the following three years due to the implemented reforms. Table 9 records the results of the three forecast tests.

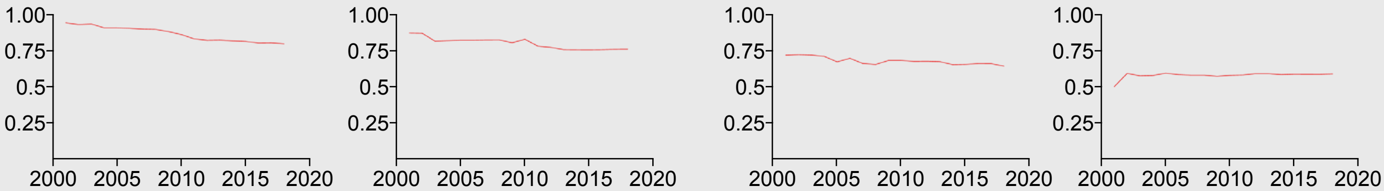
As the table shows, none of the tests indicate parameter instability in the VAR models.

Finally, we perform recursive estimation for eigenvalues, which can also be considered a valuable tool in assessing constancy in cointegrated models.[12] Figure 5 illustrates the recursively estimated eigenvalues for both VAR models.

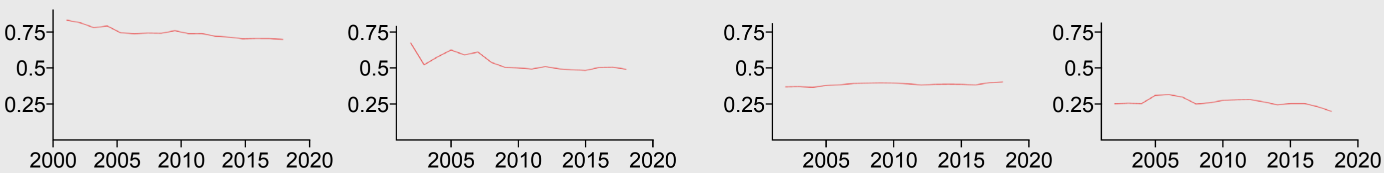
From the graphs, it can be seen that the eigenvalues from both VAR models are quite constant over the estimated period. Thus, the test results collectively suggest that the estimated parameters in the VAR models and, therefore, in the VECMs are stable over time. In other words, the results refute the claim of the nonconstancy of the estimated parameters.

Figure 5. Recursive eigenvalues.

Graph 5.A. Test results from the VAR estimated for equation (4)



Graph 5.B. Test results from the VAR estimated for equation (5)



Note: Eval = eigenvalue.

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About the Authors

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About the Projects

The KGEMM Policy and Research Studies project produces policy and applied research studies that can provide Saudi Arabian decision makers with a better understanding of domestic and international macroeconomic-energy relationships. The project mainly employs KGEMM, an energy-sector augmented general equilibrium macro-econometric model, as well as partial equilibrium frameworks.

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