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Monetary policy transmission: Balance sheet channel and investment behavior of firms in Pakistan

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Abstract

Purpose — This study investigates the relevance of the balance sheet channel of monetary policy transmission concerning non-financial firms at the Pakistan Stock Exchange (PSX), a firm-level data.

Methods — This paper estimates a family of panel data regression models and constructs a dummy variable for monetary policy tightness.

Findings — The result indicates a positive relationship between cash flows and investment during periods of monetary tightness. The impact on cash flows is visibly more pronounced than that of the quantitative effect of an increase in capital cost, which gives rise to a balance sheet channel. Three financial constraints, namely size, leverage, and dividend policy, are used to segregate firms into financially constrained and unconstrained firms.

Implication — The results highlight the balance sheet channel impact on smaller firms' cash. The cash flows of highly leveraged firms were impacted more during the tight monetary policy periods and thereby were more prone to decline in investments. Results on constraints of dividend policy are, however, inconclusive.

Originality — The paper contributes to the literature by investigating the relevance of the balance sheet channel of monetary policy transmission concerning non-financial firms using firm-level data. It also contributes to the literature by constructing a dummy variable to measure monetary policy tightness.

Keywords — Monetary policy transmission, monetary tightness, constrained firms' investments.

Introduction

Central banks use monetary policy to maintain price stability and growth. The Central bank's change in its policy rate affects all the interest rates in an economy. These changes affect the aggregate demand, which subsequently affects inflation. Economic behavior and inflation are influenced by current monetary policy and the expected future stance. Many monetary transmission channels also affect output. The significant channels are interest rate, exchange rate, asset price, credit, and expectations channels. An economy's aggregate response to the monetary policy shock will be incorporated through a combination of these channels. The effectiveness of these channels depends on the economic structure and the independence and depth of the financial sector.

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The interest rate channel affects the supply of deposits and demand for credits by influencing firm investment decisions and household spending—the credit channel supplements the interest rate channel. The credit channel affects the economy by fluctuating the supply of credit available to firms and households. The credit channel can be further subdivided into bank lending channels and balance sheet channels. According to the bank lending channel, monetary tightening influences lenders to raise lending standards due to moral hazards and adverse selection problems. On the other hand, the balance sheet channel propounds that with monetary tightening, the net-worth of collateral decreases, which leads to higher leverage and higher risk premiums, eventually restricting credit. The balance sheet channel becomes relevant because the change in monetary policy affects the financial position of firms as changes in interest rates affect the cash flows and net worth of companies.

An increase in interest rates has a twofold effect on a firm. The first is direct, i.e., interest payments on outstanding floating rate debt will increase, and secondly, the worth of the firm's collateral will decrease due to a reduction in discounted value of the firm's assets. There is an indirect consequence, i.e., due to a decrease in the net-worth of the firms. The constraint on their cash flows reduces the demand for their products, resulting in a decline in their revenues without a reduction in short-run fixed cost, which also decreases firms' net worth over time.

The purpose of this research is to examine the existence of the balance sheet channel in Pakistan. The study theorizes that external finance premium is subject to the net worth of the firms, which is because higher net worth firms have more collateral, so there is less risk for lenders. This mechanism emphasizes the role of asymmetric information in lending decisions. The informational advantage exists because the borrower is more familiar with industry dynamics and prospects (Leland & Pyle, 1977). The quality of borrowers' financial positions tends to affect the terms of their credit agreement, which means that changes in a firm's financial situation can influence their investment decisions (Dennis, Nandy, & Sharpe, 2000). This implies that investment decisions of higher net worth firms are less affected by monetary policy.

Majeed, Hashmi, and Qamar (2017) explored the effects of monetary policy on firms' fixed investment and found the relevance of interest rate and credit channels of monetary transmission in Pakistan. They concluded that the average financial conditions of the firms should be considered while devising monetary policy. The study tests whether investment decisions of manufacturing firms become increasingly sensitive to cash flows during monetary tightness based on their size and financial policy. Non-financial firms (particularly manufacturing firms) are relevant to the study because they constantly need to update their technology and facilities to grow. They require further new investments to capture markets and cater to increased demands (Brynjolfsson & Hitt, 1995; Cleary, 1999).

Previous studies have focused on the impact of monetary transmission mechanism at the aggregate level. In contrast, this study analyzes firm-level data to understand and interpret investment decisions in times of monetary tightness. The objective of the study is to see the impact of monetary policy and cashflows on firm investment. The study evaluates the effectiveness of the monetary policy. It serves as a guideline to take steps to mitigate the effects of tight monetary policy, principally for the small and medium industry.

There is a considerable debate about the effectiveness of various channels. Taylor (1995) found strong evidence for interest rate channels by analyzing the effects of interest rates on consumer spending. Bernanke and Gertler (1995) seem to contradict this assertion and state that studies have not substantiated the significance of interest rate channels and advocate the existence of credit channels. They argued that small changes in interest rates could cause an enormous impact on economic aggregates. Still, investment does not respond to these changes immediately and only fluctuates when the interest rate effect has already passed. They argued that the credit channel amplifies the impact of tight monetary policy.

The credit channel can be described as a manifestation of Lemon's problem which was coined by Akerlof (1970). The problem arises due to information asymmetry where even the firms with less risky projects might be charged a higher than the average interest rate. Bernanke and Gertler (1989) argued that due to external finance premium, the cost of external finance is always more than internal finance if the external financing is not fully collateralized. According to Stiglitz and Weiss (1981), asymmetric information may lead banks to ration loans and even deny lending to borrowers willing to pay a higher interest rate.

Mishkin (1996) pointed out a significant feature of the balance sheet channel, i.e., it is the nominal interest rates that affect a firm's cash flows and not the real interest rates. He further argued that short-term interest payments have the most significant impact on a firm's cash flows. Oliner and Rudebusch (1996) used data from 7000 manufacturing firms in the US to investigate the relationship between internal funds and investment. They concluded that monetary shocks shift all kinds of financing from small to large firms. According to them, this also results in declining aggregate bank loans because larger firms rely less on borrowing. Kashyap and Stein (2000) argued for an increase in the balance sheet channel's relevance and diminishing importance of the lending channel due to more developed financial markets and institutional changes.

Zhang and Zheng (2020) analyzed the determinants of financial investment of non-financial firms in China. They found that monetary policy is one of the significant determinants of firms' financial investments. They also concluded that the relationship between monetary policy and firms' investment is not linear and depends on the firms' nature. This study considers the heterogeneity of the firms while analyzing the impact of monetary policy on the firms' financial investment.

The same arguments have been made by Gertler and Gilchrist (1994), where they found that small firms account for a comparatively larger share of manufacturing decline during the period of monetary tightness. They attribute this decline to the size of the balance sheet of those firms, thereby indicating the existence of the balance sheet channel. Bernanke, Gertler, and Gilchrist (1996) and Bougheas, Mizen, and Yalcin (2006) found that small firms were susceptible to periods of monetary tightening compared to larger firms. Durante, Ferrando, and Vermeulen (2020) also concluded that the reaction of firms' investment to a monetary policy shock is heterogeneous. They concluded that young firms are more sensitive to the monetary policy shock than large firms. This study also considers the firms' heterogeneity while checking the impact of monetary policy transmission on firms' investment.

Kahle and Stulz (2011) suggested that US firms with more substantial balance sheets are less likely to curtail capital expenditure or change investment decisions due to monetary tightening. Crisóstomo (2012) argued that size inhibited the firm from investing in R&D. Bryson (2009) also concluded that large firms only reacted towards growth in sales, while the investment behavior of small firms is more sensitive to the cost of capital. Moreover, the balance sheet channel comes into play as smaller firms have less collateral to guarantee credits.

Ciccarelli, Maddaloni, and Peydro (2010) concluded that the bank lending channel's impact had been diminishing because of unconventional monetary policy by the US and Euro Zone while the impact of credit channel had amplified. Contrary, Łyziak, Przystupa, and Wróbel (2008) argued that reactions of various types of loans to monetary policy shocks varied greatly in Poland. The investment loans responded differently in comparison to other classes during monetary tightening. Özlü and Yalçın (2012) found similar results in Turkey. The firm size and export performance affect the composition of external finance during monetary tightening. Small firms, particularly those having fewer exports, are likely to be financially constrained (have less bank finance) during monetary tightening, so move towards trade credits aggressively, which implies that trade credit channels can subdue the traditional credit channel. In Pakistan, the prevalence of such practice cannot be overruled due to an export finance scheme by SBP that provides trade credit at subsidized rates to the exporters.

Methods

Investment is positively dependent upon cash flows of financially constrained firms, while there is no effect on unconstrained firms (Fazzari, Hubbard, & Petersen, 2000). Based on the theoretical work of Bernanke et al. (1996), Chatelain (2003), Prasetyantoko (2007), and Angelopoulou and Gibson (2009), it is expected that investment declines with monetary tightness. Further, it is also likely that the decline is more pronounced for financially constrained firms as their cash flows inhibit them from the investment during such periods.

Following Fazzari, Hubbard, and Petersen (1988), the following equation is used in the study, which has already been used by Prasetyantoko (2007) and Crisóstomo (2012).

$$Gross\ Investment_{it} = \beta_0 + \beta_1(Average\ Q_{it}) + \beta_2(Cash\ Flows_{it}) + \beta_3(Cash\ Flows * MPDummy_{it-1}) + \varepsilon_{it}$$

The study uses secondary data of 217 manufacturing firms registered in PSX from the period 1999 – 2015. The manufacturing firms must regularly update their technology and facilities to grow, so their investment decisions are more sensitive to cash flows. The financial statement data are obtained from reports published by the State Bank of Pakistan, whereas stock market information is obtained from PSX.

Gross investment (investment behavior) is computed by calculating the difference between the book value of assets in two consecutive periods (Fazzari et al., 2000). The difference is normalized by total assets at replacement cost averaged across period's *t*-1 and *t*.

A dummy variable for monetary policy tightness is constructed, where it is measured using a change in policy rates. For identifying monetary policy episodes, monetary policy statements (MPS) of all periods from 1999 to 2015 were taken. Based on the above characteristics, four periods have been identified as policy episodes in the period under review: (i) 1999 – Oct 2001, (ii) Oct 2001 – July 2007, (iii), July 2007 – Oct 2014, and (iv) Oct 2014 – 2015.

A dummy is constructed for the years during which monetary policy stance is analyzed and given a value of 1 in case of a tight episode of monetary policy and a value of 0 otherwise. However, the monetary policy dummy is monthly, and firm-level data is annual for Cash flows and gross investment. Therefore observations are averaged across the whole financial year and will be given the value of 0 or 1 based on the number of months (Gertler & Gilchrist, 1994; Rondi, 1998; Haan & Sterken, 2000; Jimenez, 2012). Finally, the cash flows are calculated by the ratio of cash flow from operations minus interest payments and total assets at replacement cost averaged over periods t and t-1. (Fazzari et al., 1988; Lang, Ofek, & Stulz, 1996).

Average Q (market value of a firm to the replacement cost of assets) is the firm's stock market value divided by total assets at replacement cost averaged over periods *t,* and *t-1* is average Q.

Constrained Firm and Unconstrained Firms

The study will use three classification criteria and apply descriptive statistics to distinguish financially constrained firms from others. Firstly, the sample is divided into large and small firms based on net worth. The firms that lie in the bottom 33% quartile will be considered small. Secondly, companies are classified as financially constrained based on dividend announcement, where non-paying firms are classified in the financially constrained category. The third criteria use the leverage ratio, where the firms with relatively high leverage ratios are expected to be financial.

Variables Measurements

The variables are measured hereunder:
Average Q = Market Value/Total Assets at Replacement Cost
Cash Flows = (Cash flows from Operations – Interest Payments)/Total Assets at Replacement cost
CF*MPD = Cash Flows under Monetary Tightness

Results and Discussion

Table 1 shows the descriptive analysis of the variables under the tight monetary policy. The average gross investment in fixed capital is 3.5%, even without discounting for tighter monetary policy. The median value shows that more than half of the companies have positive investment figures. The standard deviation of the data indicates that without accounting for periods or size, the investment rate of firms has a low degree of variation. Tobin's Q means on the lower side (between

0 and 1), which implies that the cost to replace a firm's assets is greater than the value of its stock. On the other hand, cash flows are more than Gross investment, i.e., 6.8%, which means that all excess cash flows are not utilized in investment as some portion goes into retained earnings or dividends. Further, the cash flows and cash flows under tight monetary policy relationship is observable as cash flows are nearly half during periods of monetary tightness.

Table 1. Descriptive Statistics

	Gross Investment	Average Q	Cash Flows	CF*MPD
Mean	0.036	0.338	0.068	0.034
Median	0.026	0.197	0.046	0.000
Std. Dev.	0.196	0.411	0.138	0.081
Skewness	-1.803	3.133	-6.355	2.668
Kurtosis	22.025	16.795	200.873	24.391
Observations	3472	3472	3472	3472

Correlation Matrix

In table 2, the correlation coefficients of the variables are provided, the matrix shows the low correlation among the variables, so there is no chance for multicollinearity. Furthermore, Average Q and Cash Flows are positively correlated with Gross Investment, whereas cash flows under the tight monetary policy are negatively correlated with Gross Investment.

Table 2. Correlation Matrix

	Gross Investment	Average Q	Cash Flows	CF*MPD
Gross Investment	1			
Average Q	0.045	1		
Cash Flows	0.074	0.269	1	
CF*MPD	-0.001	0.076	0.481	1

Regression Results

The estimation of the basic model, which is our first empirical exercise, is presented in Table 3 below.

Table 3. Fixed Effect Model Results

Variable	Coefficient	Std. Error	t-Statistics	Prob.
C	0.023	0.006	4.115	0.000***
Average Q	0.028	0.012	2.317	0.021**
Cash Flows	0.1085	0.031	3.499	0.001***
CF*MPD	-0.127	0.048	-2.645	0.008***
		Weighted Statistics		
R-squared	0.073		Adjusted R-squared	0.011
F-statistic	1.170		Prob(F-statistic)	0.049**
Note: ***, **, * denote the level of significance at 1%, 5% and 10% respectively.				
Number of Observations: 3472		Period: 2000 – 2015		

The Average Q is positive and statistically significant at a 5% level of confidence, suggesting that investment decisions are sensitive to investment opportunities. The Q value is a significant determinant of investment. The cash flows are also positive and statistically significant at a 1% level of confidence, which implies that companies with more internal resources tend to invest more than firms with limited internal resources. These results align with Myers and Majluf (1984) and Kashyap, Stein, and Wilcox (1993). The cash flows under the tight monetary policy are negatively related to gross investment, which implies that during periods of monetary tightness, the firm's cash flows are reduced, thereby requiring external financing to maintain the level of investment. In times of monetary tightness, financial constraints can increase significantly; however, it cannot be assumed that it would affect the relationship between cash flows and investment. The balance sheet channel can be observed from the above relationship, thereby establishing a negative relationship between cash flows and investment during monetary tightness. To check the sensitivity of cashflows to investment, the sample is split using three criteria i.e., size, dividend policy and leverage.

Regression Results: Firms Constrained by Size

When firms are separated based on their size and separate regression is run based on constrained and unconstrained firms, the results signify a balance sheet channel. However, products of un-constrained firms (Table 4) are somewhat like those of constrained firms concerning coefficients; however, the Average Q's significance concerning investment falls below 10%. For unconstrained firms' same level of investment can be arranged to require a lower level of cash flows. The impact of tight monetary policy is visible on investment in unconstrained firms, which is quite like the aggregate data. In constrained firms (Table 5), the average Q is insignificant. However, the coefficient of cash flows under tight monetary policy reduces significantly from -10.8% in unconstrained firms to -26.3% in constrained firms.

Table 4. Fixed Effect Model Results

Variable	Coefficient	Std. Error	t-Statistics	Prob.
C	0.012	0.008	1.521	0.128
Average Q	0.029	0.018	1.595	0.111
Cash Flows	0.087	0.038	2.307	0.021**
CF*MPD	-0.109	0.061	-1.792	0.073*
	Weighted Statistics			
R-squared	0.063		Adjusted R-squared	-0.006
F-statistic	0.920		Prob(F-statistic)	0.746
	Hausman specification test			
Chi-Sq. Statistic	21.851		Prob	0.000***

Table 5. Random Effect Model Results

Variable	Coefficient	Std. Error	t-Statistics	Prob.
C	0.060	0.008	7.897	0.000***
Average Q	0.008	0.011	0.676	0.500
Cash Flows	0.116	0.050	2.350	0.019**
CF*MPD	-0.264	0.069	-3.828	0.000***
		Weighted Statistics		
R-squared	0.015		Adjusted R-squared	0.012
F-statistic	5.424		Prob(F-statistic)	0.001***
		Hausman specification test		
Chi-Sq. Statistic	5.364		Prob	0.147

Note: ***, **, * denote the level of significance at 1%, 5% and 10% respectively.
Number of Observations: 1072 Period: 2000 – 2015

The results are consistent with our hypothesis as smaller firms' cash flows are more constrained than larger firms requiring more financing from external resources. This explains the two-fold impact of tight monetary policy, i.e., the smaller firms are more impacted by tight monetary policy than larger firms in terms of both access to finance and cost of finance and will need to put up more collateral to finance investments. The results corroborate the findings of Gertler and Gilchrist (1994), Bougheas et al. (2006), and Crisóstomo (2012). The results also support the hierarchy of finance theory as pronounced by Oliner and Rudebusch (1992). It is concluded that in times of monetary tightness, the cash flows of a firm become an important determinant of investment, particularly for smaller firms. On the one hand, the policy reduces the company's cash flows as a result of increasing cost of capital, while on the other side net worth of the firms decline making it harder for them to raise external finance.

Regression Results: Firms Constrained by Leverage

Tables 6 and 7 show regression results for firms unconstrained and constrained by leverage. The results of unconstrained firms are again similar to those of the main regression equation. Unconstrained firms have free cash flows to benefit from investment opportunities available to them; however, the results are not significant. Average Q is not a significant explanatory variable for either constrained or unconstrained firms as the results are insignificant for both and even more for constrained firms. Cash flows explain the level of investment reasonably well for both classes as the results are significant. The results on firms' cash flows constrained by leverage in times of tight monetary policy are even more striking (Table 7). Firms constrained by the leverage (Highly leveraged) require even higher cash flows during regular periods to finance investment, i.e., 45.9% (Table 7) compared to 9.6% (Table 6). This is mainly because these firms have a lesser ability to raise more external finance. In periods of monetary tightness, the impact of leverage on cash flows seems much pronounced, i.e., the cash flows coefficient reduces from -10.3% to -38.8%. This implies that it becomes even harder for firms constrained by leverage to raise external finance, particularly in periods of monetary tightness. During tight monetary policy, a firm is already constrained by falling sales revenues and increasing cost of capital, where falling revenues also reduces the value of collateral, thereby reducing the firm's ability to carry out investment, propagating the existence of a balance sheet channel. It becomes even hard for the firm to maintain the production capacity of its existing machinery due to a lack of finance. These results support the findings of Lang et al. (1996) and Gedajlovic (2005) and conclude that investment by firms financially-constrained by leverage suffers more during periods of monetary tightness.

Table 6. Random Effect Model Results

Variable	Coefficient	Std. Error	t-Statistics	Prob.
C	0.033	0.006	5.862	0.000***
Average Q	0.012	0.009	1.333	0.183
Cash Flows	0.097	0.030	3.237	0.001***
CF*MPD	-0.104	0.050	-2.086	0.037**
		Weighted Statistics		
R-squared	0.006		Adjusted R-squared	0.005
F-statistic	5.217		Prob(F-statistic)	0.002***
		Hausman specification test		
Chi-Sq. Statistic	7.449		Prob	0.059

Note: ***, **, * denote the level of significance at 1%, 5% and 10% respectively.
Number of Observations: 2608 Period: 2000 – 2015

Table 7. Random Effect Model Results

Variable	Coefficient	Std. Error	t-Statistics	Prob.
C	0.017	0.008	2.071	0.039**
Average Q	-0.017	0.030	-0.574	0.566
Cash Flows	0.455	0.110	4.192	0.000***
CF*MPD	-0.388	0.153	-2.530	0.012**
		Weighted Statistics		
R-squared	0.021		Adjusted R-squared	0.017
F-statistic	6.010		Prob(F-statistic)	0.001***
		Hausman specification test		
Chi-Sq. Statistic	7.045		Prob	0.071

Note: ***, **, * denote the level of significance at 1%, 5% and 10% respectively.
Number of Observations: 864 Period: 2000 – 2015

Regression Results: Firms Constrained by Dividend Policy

Tables 8 and 9 show regression results of unconstrained firms and firms constrained by dividends. The results of unconstrained firms are in line with the normal regression equation, with all coefficients being significant. However, the results of constrained firms are more interesting as they do not correspond to the results of either the main equation or other constrain variables (size and leverage). Average Q has a negative relationship with gross investment, while cash flows under the tight monetary policy have a positive relationship. Although the results of Average Q are significant, the results are not significant for cash flows under the tight monetary policy. The results of unconstrained-firms support existing literature; however, the results of constrained-firms do not correspond with the literature. This could be due to several reasons, i.e., the data is limited in the sense as we have only ascribed those firms as constrained that have not given dividends throughout the period. At the same time, firms that announce dividends one year may have remained constrained during the rest of the periods. Further, the number of observations might below have results as per the existing empirical evidence.

Table 8. Fixed Effect Model Results

Variable	Coefficient	Std. Error	t-Statistics	Prob.
C	0.028	0.006	4.917	0.000***
Average Q	0.032	0.012	2.723	0.007***
Cash Flows	0.101	0.030	3.356	0.001***
CF*MPD	-0.142	0.047	-3.021	0.001***
		Weighted Statistics		
R-squared	0.067		Adjusted R-squared	0.004
F-statistic	1.066		Prob(F-statistic)	0. 257
		Hausman specification test		
Chi-Sq. Statistic	10.535		Prob	0.015**
Note: ***, **, * denote the level of significance at 1%, 5% and 10% respectively.				
Number of Observations: 3168			Period: 2000 – 2015	

Table 9. Random Effect Model Results

Variable	Coefficient	Std. Error	t-Statistics	Prob.
C	0.002	0.023	0.076	0.940
Average Q	-0.158	0.073	-2.168	0.031**
Cash Flows	0.396	0.187	2.120	0.035**
CF*MPD	0.065	0.247	0.262	0.794
		Weighted Statistics		
R-squared	0.033		Adjusted R-squared	0.022
F-statistic	3.142		Prob(F-statistic)	0.026**
		Hausman specification test		
Chi-Sq. Statistic	4.044		Prob	0.257
Note: ***, **, * denote the level of significance at 1%, 5% and 10% respectively.				
Number of Observations: 285			Period: 2000 – 2015	

Some papers such as Oliner and Rudebusch (1992), Myers and Majluf (1984), and Cleary (1999) have enforced the idea that firms prefer internally generated financing for investments which builds an argument that a reduction in investments during monetary tightness is not necessarily related to increases in the cost of capital, as firms can continue to keep investing using internal sources. However, many studies have found a definite decline in firms' investments during periods of tight monetary policy (Fazzari et. al., 1988; Bougheas et al., 2006; and Bryson, 2009). This impact of monetary policy on investment is visible and persistent across sectors and firms of different sizes. However, this study mostly focused on the balance sheet channel's existence and how it amplifies the impact of tight monetary policy on firms based on different financial constraints.

The study results indicate a positive relationship between cash flows and investment, which is in line with study of Kaplan and Zingales (1997). The results further indicate that during periods of monetary tightness, the cash flows available to the firm reduce significantly, thereby having an impact on investments of the firm. The impact on cash flows is visibly more pronounced than the quantitative impact of an increase in the cost of capital, which gives rise to the notion of the balance sheet channel's existence as theorized. The results are in line with the findings of Zhang and Zheng (2020). However, the endogeneity of the quantity of credit makes it difficult to establish that only balance sheets are channeled to those. To further substantiate the results, three financial constraints, i.e., size, leverage, and dividend policy, are used to segregate firms into financially constrained and unconstrained firms. It is assumed that constrained firms will show a stronger reaction to tight monetary policy than unconstrained firms. The balance sheet of these firms inhibits them from raising finance externally or even internally, thereby exerting pressure on their investments. Even in the presence of viable investment opportunities, the investors may be unwilling to place their money in these firms (Bernanke et al. (1996) due to the firm's inability to provide adequate collateral. The size of the firm may be a better proxy (Fazzari et al., 1988; Crisóstomo, 2012) to assess the balance sheet channel's impact as it directly correlates to a firms' ability to arrange collateral to generate funds for investments. The results highlight the balance sheet channel's existence as the impact on smaller firms' cash flows, and thereby investments are much more pronounced compared to larger firms. The results of firms constrained by leverage also show the same conclusion as cashflows of highly leveraged firms is impacted more during tight monetary periods, so these firms are more prone to decline in investments than unconstrained firms. These results are in line with the findings of Lang et al. (1996) and support a balance sheet channel. The different results based on a different level of constraints corroborate Durante et al. (2000).

Conclusion

This study has investigated the relevance of the balance sheet channel of monetary policy transmission concerning Pakistan using firm-level data. Further, normative indicators have been used to identify sustained periods of tight monetary policy in Pakistan, which has not been attempted previously in Pakistan. The research has also tried to contribute to the literature on financial constraints and investment decisions by integrating firm-level data with macro-economic information and analyze the linkage between investment and finance. The balance sheet channel is considered relevant mainly because firms are constrained by credit channel or interest rate channel as most studies have advocated. A balance sheet channel exists because credit is not provided without collateral. The value of such collateral is directly related to the amount of financing available to a firm. Banks are only willing to finance the percentage of collateral provided by a firm. During periods of tight monetary policy, the value of a firm reduces due to several factors, i.e., a decrease in aggregate consumption resulting in declining sales, increasing cost of sales due to increased costs (not just financing cost but also increase in the price of raw materials).

Finally, the dividend policy of the firm has also been used as a financial constraint. The results are, however, inconclusive concerning the Pakistani firms. However, the decline in cash flows of constrained firms is visible; however, the results are not significant. This may be due to very few observations for firms that were constrained by the inability to give dividends. Another reason could be that firms are classified as unconstrained even if they paid a dividend in only one period, which may have led to inconclusive results. However, evidence for the balance sheet channel's existence is substantial considering the results of other financial constraints, i.e., size and leverage.

The study has some strong implications for the policymakers to identify if rate subsidies (to support exports) have any positive impact during periods of monetary tightness. SBP also offers rate subsidies on long-term financing facilities, which could also be studied in conjunction with other rate subsidies. The existence of cash flows and investment relationships suggests that corporate managers must consider cash flows while devising the investment strategies. The monetary transmission channel's impact must also be studied to develop strategies for the investment behavior, as the transmission channels play a key role in the change in behavior of cashflows and investment.

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The role of agricultural productivity in economic growth in middle-income countries: An empirical investigation

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Abstract

Purpose — This study investigates the role of agricultural productivity in economic growth in middle-income countries.

Methods — This study utilizes the data of 53 middle-income countries over the period 1991-2017 and provides robust estimations using second-generation panel data methods considering cross-sectional dependency.

Findings — The estimation results of the Common Correlated Effects Mean Group (CCEMG), Dynamic-CCEMG, and biased-corrected form of Dynamic-CCEMG, suggest that agricultural productivity is the main engine of economic growth. Additional findings show that economic growth is positively associated with both physical capital and human capital. This paper does not find any significant relationship between trade openness and economic growth.

Implications — This study reveals that the industrialization process in middle-income countries to boost economic growth can be accelerated by implementing policies to increase productivity in the agricultural sector.

Originality — This study focuses on analyzing the effect of agricultural productivity neglected mainly in recent studies on economic growth. This paper develops a second-generation estimator that considers cross-sectional dependence.

Keywords — Economic growth, agricultural productivity, human capital, trade openness.

Introduction

The early literature of economic development has shown modern industrialization as the basic dynamic of economic growth, while the agricultural labor force was a source of cheap labor to feed industrialization. Agriculture is perhaps the only sector that produces output in the countries at the early stages of development, and in these countries, the majority of society is employed in the agricultural sector (Johnston & Mellor, 1961). In the period when the economy is transforming, there are opinions that the workers in the agricultural sector are the primary source of industrial development. In his study, Lewis (1954) claimed that while the industrial sector was growing, the agricultural sector would be a source of labor with sufficiently low wages in his two-sector model. With this "unlimited" supply of labor, agriculture would be a source of capital formation because of accessing cheap labor and yet, lower production costs. He showed the hidden unemployment in the agricultural sector as a cause of this sectoral transition. According to the World Food and

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Agriculture Statistical Pocketbook published by FAO (Food and Agriculture Organization of the United Nations), the percentage share of agricultural value-added diminishes with the development of a country (FAO, 2018). This report supports the idea proposed by Arthur Lewis. With the development of industrial sectors, the share of agriculture in GDP goes down because of the migration of workers from agriculture to the industrial sector. In the report of FAO, it is also mentioned that in order to keep agriculture moving, agricultural productivity per-worker should be increased.

At the end of the twentieth century, the dominant view on development economics suggested that the share of agriculture in the economy would naturally diminish, and therefore there was no need for modernization in agriculture. While Lewis's (1954)model disregarded efforts to modernize agriculture, Hirschman's (1958) approach prioritized the development of the manufacturing industry due to linkage effects (Timmer, 2002). Besides, Prebisch (1950) argued that the increase in agricultural production would cause deterioration in terms of trade. These approaches, which emerged in development theories, have resulted in not paying the necessary attention to the modernization of agriculture and productivity increase. Although the share of agriculture in the overall economy is diminishing, modernization of agriculture and increasing agricultural productivity are important for achieving more output by fewer workers in the agricultural sector. There are significant differences between high, middle, and low-income countries in the context of per worker output in the agricultural sector. To illustrate this difference, a comparison is given in Table 1.

Table 1. A comparison of agricultural productivity between different income groups (2016)

	Agricultural value-added per worker (constant 2010)	Agricultural value-added (% gdp)
High-income	36 799 $	1.3
Middle-income	2990$	15.7
Low-income	599 $	25.4

Source: World Bank, World Development Indicators (2019).

According to Table 1, there is evidence supporting the view that the share of agriculture in the economy decreases as the level of income increases. However, the high-income countries' agricultural output per worker is much higher than middle and low-income countries. In low-income countries, the agricultural sector accounts for 25% of national income, while in middle-income countries, 15.7% and high-income countries only 1.4%. According to World Development Indicators (2019), middle-income countries produce 74% of world agricultural output. However, there are significant productivity differences between high and middle-income countries. An agricultural worker in a high-income country is about 12.3 times more productive than in middle-income countries. Such a productivity difference may be a result of the effort to increase productivity. According to a global assessment of world agricultural R&D spending published by Beintema, Stads, Fuglie, and Heisey (2012), 51% of public agricultural R&D investments are made in high-income countries. Also, most of the private sector R&D investments are made by companies in developed OECD countries. Despite the neoclassical point of view, efforts to modernize agriculture would contribute to economic growth by increasing productivity. These efforts are also important to provide enough food for the rising human population worldwide. Increasing the need for food together with the ever-increasing population requires a continuous increase in agricultural productivity (Mariyono, Kompas, & Grafton, 2010).

In the context of agriculture's traditional role in economic growth, Johnston and Mellor (1961) suggested four more linkages in addition to agriculture's role in providing cheap labor to industrial sectors. The first is that the agricultural sector provides food for consumption. Rising domestic food supply provides high-calorie nutrition to the workers. Since production in developing countries is predominantly based on unskilled labor, continuity of food production is important in meeting the food needs of workers. It also leads to stability in food prices and facilitates access to food (Timmer, 1995). The second linkage they established is from agricultural exports to economic growth. In the early stages of development, agricultural exports can be an important instrument to increase foreign exchange earnings and revenue. Countries that have achieved comparative advantages with industrial production will prefer to import agricultural food products (Nicholls, 1964). Less developed countries do not have enough capital accumulation to have a comparative advantage over industrial products. Therefore, agricultural exports are the main sources of foreign exchange earnings. The third is the contribution to investments. Since most of the total production in underdeveloped countries is carried out in the agricultural sector, the development of the industrial sector depends on the contribution of the agricultural sector to capital formation. The fourth is that increasing incomes in the agricultural sector can increase the demand for industrial products and contribute to the expansion of the industrial sector with a wider market.

Besides linkages proposed by Johnston and Mellor (1961), Kuznets (1964) and Ghatak and Ingersent (1984) explained the contribution of agriculture in economic growth in four different ways that are similar to the Johnston and Mellor's but provide a classification. These four ways are product contribution, market contribution, factor contribution, and foreign exchange contribution. According to Khan, Jamshed, and Fatima (2020), product contribution refers to the growth of product per capita, while market contribution represents the rising demand for industrial inputs such as fertilizer, pesticide, farm machinery, etc. Factor contribution refers to the shift of labor in the agricultural sector to non-agricultural sectors. Since the increase in productivity resulting from the increase in the use of technology in agriculture will provide higher output with less labor and capital use, these factors are employed in industry instead of agriculture. Finally, foreign exchange contribution is the contribution of agricultural exports to countries' foreign exchange earnings. In addition to its positive impact, agriculture may retard economic growth. Educational demands of agricultural workers are generally lower than those in trade, industry, and service sectors (Gylfason, 2000). Increasing the share of agriculture in the economy may lead to a decrease in overall human capital formation. Besides, the allocation of the labor force to the industrial sector may be higher in an economy with a less productive agricultural sector (Matsuyama, 1992). The allocation of more resources to the industrial sector can lead to faster economic growth.

In addition to theoretical discussions, there are several empirical studies in the previous literature. Hwa (1988) provided evidence of the contribution of agriculture to economic growth. Cross-sectional data of 87 countries covering the period 1960s and 1970s were used in this study. Cross-section regression results show that agricultural growth has a significant positive effect on overall economic growth. Yao (2000) concluded that agriculture is one of the driving forces of Chinese economic growth, supporting Hwa's (1988) evidence. The author also argued that biased policies against agriculture weaken the contribution of agriculture to economic growth. Irz, Lin, Thirtle, and Wiggins (2001) found a significant effect of agricultural productivity growth on rural poverty and human development. According to the results, rising agricultural productivity reduces poverty and increases human development. Gollin, Parente, and Rogerson (2002) provided panel data evidence using the data from 62 developing countries. The author concluded that agricultural productivity has a positive impact on economic growth. The movement of labor from agriculture to non-agricultural sectors also improves economic growth. According to this result, the allocation of the labor force to the industrial sector with an increase in per-worker agricultural productivity is effective in the economic development process. As opposed to this, Gardner (2005) found no causality between agriculture and overall economic growth. Following this study, Tiffin and Irz (2006) investigated the relationship between agricultural value-added and economic growth with a panel of 85 countries.

According to Granger causality results, there is unidirectional causality from agriculture to economic growth. Another study investigating the causality between agriculture and economic growth is proposed by Katırcıoğlu (2006). He used annual data for North Cyprus in the period 1975-2002. The results indicate that there is bidirectional causality between agriculture and economic growth. While some studies have investigated the direct impact of agriculture on growth, while others have investigated the impact on non-agricultural sectors. Timmer (2002) found a significant positive impact of agricultural GDP growth on non-agricultural GDP growth. In his study, a panel of 70 countries was used, covering the five-year averages of the period 1960-1985. Bravo-Ortega and Lederman (2005) revisited the analysis of Timmer (2002) and concluded that was supporting Timmer's study. They used a wider panel data set than Timmer, covering five-year averages 1960-2000 with different agricultural growth definitions.

According to pooled OLS and fixed effects estimation results, non-agricultural growth is positively associated with agricultural growth. Chebbi (2010) examined the cointegration and causality between the agricultural and non-agricultural sectors in Tunisia. Cointegration results show a long-run relationship between agricultural GDP, real GDP of manufacturing and non-manufacturing industries, tourism and transportation, and commerce and services sectors. Also, agricultural growth causes growth in the agro-food industry. Oyakhilomen and Zibah (2014) found a significant positive relationship between agricultural output and economic growth in Nigeria utilizing time-series data covering 1970-2011. Awokuse and Xie (2015) also examined the cointegration and causality between agriculture and economic growth. They used annual data for 9 developing and transition economies covering the period 1980-2011. The results obtained from the ARDL model established to analyze cointegration show a significant long-run relationship between agriculture and economic growth for all countries included in the model. Contemporaneous causality results obtained from dynamic acyclic graphs (DAGs) show a significant causality from GDP to agriculture in Chile, Mexico, Kenya, and South Africa. The causality from agriculture to GDP is significant in China, Indonesia, Thailand, and Cameroon. In Brazil, no significant causal relationship was found in their study. Sertoglu, Ugural, and Bekun (2017) also investigated the agriculture-economic growth nexus in Nigeria. Their time-series analysis is covering the period 1981-2013. They came to the same conclusion as Oyakhilomen and Zibah (2014). In their recent study, Khan et al. (2020) found that the agricultural sector positively impacts other sectors' economic growth in West Bengal. They concluded a long-run causality between the agricultural and industrial sectors, the service sector, and overall economic growth.

This study aims to investigate the role of agricultural productivity in economic growth in middle-income countries. Agriculture has an important place in the production and employment of middle-income countries. However, these countries are behind high-income countries in terms of agricultural productivity. According to Mariyono (2019), the widespread use of technology in agriculture and increased productivity are important for economic growth and rural development in emerging economies. However, it is observed that the effects of agricultural productivity are mostly neglected in recent studies on economic growth. In addition to previous empirical researches, this study provides second-generation panel data estimations. The empirical works in previous literature consist of traditional estimation methods called first-generation tests. In the analysis of panel data, the estimators considering cross-sectional dependence are called the second-generation estimators. Cross-sectional dependency simply refers to the situation when the shock that occurs in one country affects other countries as well. The source of this problem encountered in panel data analysis is the economic, financial, and political integration among countries (Menyah, Nazlioglu, & Wolde-Rufael, 2014).

The assumption of cross-sectional independence is hard to satisfy because of the high level of economic integration among countries. İgnoring this effect may lead to misleading inferences (Chudik & Pesaran, 2013; De Hoyos & Sarafidis, 2006). According to Awokuse and Xie(2015) and Tsakok and Gardner (2007), previous studies in the agricultural sector's contribution to economic growth do not provide a consensus due to methodologically weak empirical analysis. Therefore, fresh analyzes are required. As mentioned above, middle-income countries have the largest share in the world's agricultural production than other income groups. Especially in those countries that are in a transformation from agriculture to industry and the service sector, it is important to reveal the contribution of the development in the agricultural sector to economic growth in order to shape the policies to be implemented. In our study, empirically robust results with new panel data analysis methods are provided. Besides, since economic relations are often in a dynamic adjustment process (Bun & Sarafidis, 2015), dynamic panel data estimators were used.

Methods

In empirical studies in economic growth, economic growth is explained mainly by the expansion of the physical and human capital formation (Barro, 1991, 2001; Mankiw, Romer, & Weil, 1992; Romer, 1989; Sala-i-Martin, 1997). Some studies provide evidence that trade openness has a positive impact on economic growth (Edwards, 1998; Frankel & Romer, 1996; Sachs, Warner, Åslund, & Fischer, 1995). Following the related literature, the estimated model in this study as follows;

$$GDP_{it} = \beta_{1i} + \beta_2 AGR_{it} + \beta_3 GCF_{it} + \beta_4 HC_{it} + TRD_{it} + \varepsilon_{it} \quad (1)$$

In equation 1, GDP is gross domestic product per capita (constant 2010 \$), GCF is gross capital formation percentage share of GDP, HC is human capital indicates average years of schooling based on the estimations of Barro and Lee (2013) and a rate of return to education based on the estimations of Psacharopoulos (1994). TRD is trade openness which refers to the sum of total exports and total imports percentage share of GDP. In the equation, AGR indicates agricultural productivity. Following Gardner (2005) and Tiffin and Irz (2006), agricultural value-added per worker (constant 2010 \$) was used as an indicator of agricultural labor productivity.

GDP, GCF, TRD, and AGR data are obtained from World Bank World Development Indicators (2019) while HC data from Penn World Tables (Feenstra, Inklaar, & Timmer, 2015). Our dataset consists of data from 53 middle-income countries.[1] The period is covering 1991-2017. All variables are turned into a logarithmic form. In the estimation of the model, second-generation panel data analysis methods were used.

Cross-sectional Dependence and Slope Homogeneity

Standard panel data analysis methods assume that no dependency exists between cross-section units and slope coefficients are homogenous. However, the estimators ignoring the cross-sectional dependence can result in misleading inference (Chudik & Pesaran, 2013). In addition, the estimated coefficients may differ across cross-section units. Therefore, the existence of cross-sectional dependence and slope homogeneity will be investigated at first. The existence of cross-sectional dependence in the error term obtained from the model analyzed with Pesaran (2004) CD_{LM} and Pesaran, Ullah, and Yamagata (2008) bias-adjusted *LM* test. These methods are valid while N>T and T>N. However, according to Pesaran (2004), there would be size distortions when N is relatively large. In our study, the time dimension is 27, and the cross-section dimension is 53 (N>T). Therefore, both CD_{LM} and bias-adjusted *LM* (LM_{adj}) tests are used. Their test statistics can be calculated as follows:

$$CD_{LM} = \left(\frac{1}{N(N-1)}\right)^{\frac{1}{2}} \sum_{i=1}^{N-1} \sum_{j=i+1}^{N} (T\hat{\rho}_{ij}{}^2 - 1) \quad (2)$$

$$LM_{adj} = \sqrt{\frac{2}{N(N-1)}} \sum_{i=1}^{N-1} \sum_{j=i+1}^{N} \frac{(T-k)\hat{\rho}_{ij}^2 - \mu_{Tij}}{v_{Tij}} \quad (3)$$

Equation 2 shows the calculation of Pesaran (2004) CD_{LM} and equation 3 is Pesaran et al. (2008) bias-adjusted *LM* test statistic. V_{Tij}, μ_{Tij}, and $\hat{\rho}_{ij}$ respectively represent variance, mean, and the correlation between cross-section units. The null and alternative hypothesis for both test statistics;

H_0: No cross-sectional dependence exists

H_1: Cross-sectional dependence exists

Pesaran and Yamagata (2008) developed Swamy's (1970) random coefficient model to investigate parameter heterogeneity in panel data analysis.

Swamy's test statistic can be calculated as follows.

[1] Those countries are Albania, Argentina, Bangladesh, Belize, Bolivia, Botswana, Brazil, Bulgaria, Cameroon, China, Colombia, Congo, Costa Rica, Dominican Republic, Ecuador, Egypt, El Salvador, Eswatini, Gabon, Guatemala, Honduras, India, Indonesia, Iran, Jamaica, Jordan, Kenya, Kyrgyz Republic, Malaysia, Mauritania, Mauritius, Mexico, Mongolia, Morocco, Namibia, Nicaragua, Nigeria, Pakistan, Paraguay, Peru, Philippines, Romania, Russian Federation, Senegal, South Africa, Sri Lanka, Sudan, Thailand, Tunisia, Turkey, Ukraine, Vietnam, and Zimbabwe.

$\hat{S} = \sum_{i=1}^{N}\left(\tilde{\beta}_i - \tilde{\beta}_{WFE}\right)\frac{x_i' M_T x_i}{\widetilde{\sigma_i^2}}\left(\tilde{\beta}_i - \tilde{\beta}_{WFE}\right)$ (4)

In equation 4, $\tilde{\beta}_i$ and $\tilde{\beta}_{WFE}$ respectively indicate the parameters obtained from pooled OLS and weighted fixed effects estimation while M_T is the identity matrix. Swamy's test statistic is developed by Pesaran et al. (2008) with the following equations,

$$\tilde{\Delta} = \sqrt{N}\left(\frac{N^{-1}\tilde{S} - k}{\sqrt{2k}}\right) \quad (5)$$

$$\tilde{\Delta}_{adj} = \sqrt{N}\left(\frac{N^{-1}\tilde{S} - E(\tilde{Z}_{it})}{\sqrt{Var(\tilde{Z}_{it})}}\right) \quad (6)$$

where $\tilde{S}$ is the Swamy test statistic and k is the number of explanatory variables. $\tilde{\Delta}_{adj}$ is a bias-adjusted version of $\tilde{\Delta}$. $\tilde{Z}_{it}$=k and $Var(\tilde{Z}_{it}) = 2k(T - k - 1)/T + 1$. The null and alternative hypothesis for both test statistics is given below.
$H_0: \beta_i = \beta$
$H_1: \beta_i \neq \beta$

The rejection of the null hypothesis shows the heterogeneity of slope coefficients in panel data models. After these preliminary analyses, stationarity levels of the variables will be examined with Pesaran's (2007) Cross-sectionally Augmented Dickey-Fuller (CADF) test.

Unit Root Test

Pesaran (2006) suggested a factor modeling approach that adds the cross-section averages as a proxy of unobserved common factors into the model to prevent the problems caused by cross-sectional dependence. Following this approach, Pesaran (2007) proposed a unit root test. This method is based on augmenting the Augmented Dickey-Fuller (ADF) regression with lagged cross-sectional mean and its first difference to deal with cross-sectional dependence (Baltagi, 2008). This method considers the cross-sectional dependence and can be used while N>T and T>N. The CADF regression is;

$$\Delta y_{it} = \alpha_i + \rho_i^* y_{i,t-1} + d_0\bar{y}_{t-1} + d_1\Delta\bar{y}_t + \epsilon_{it} \quad (7)$$

$\bar{y}_t$ is the average of all N observations. To prevent serial correlation, the regression must be augmented with lagged first differences of both y_{it} and $\bar{y}_t$ as follows;

$$\Delta y_{it} = \alpha_i + \rho_i^* y_{i,t-1} + d_0\bar{y}_{t-1} + \sum_{j=0}^{p} d_{j+1}\Delta\bar{y}_{t-j} + \sum_{k=1}^{p} c_k\Delta y_{i,t-k} + \epsilon_{it} \quad (8)$$

After this, Pesaran (2007) averages the t statistics of each cross-section unit ($CADF_i$) in the panel and calculates *CIPS* statistic as follows;

$$CIPS = \frac{1}{N}\sum_{i=1}^{N} CADF_i \quad (9)$$

The null hypothesis of this test is the existence of a unit root in the panel in question. After the preliminary analysis of unit root, the long-run relationship will be investigated via Westerlund's (2008) Durbin-Hausman cointegration test and Dynamic CCEMG (Dynamic Common Correlated Effects Mean Group) estimator developed by Chudik and Pesaran (2015).

Cointegration and Long-Run Estimation

In our study, we investigate the cointegration relationship between the variables using the Durbin-Hausman test developed by Westerlund (2008). Cross-section dependence is allowed by modeling common factors by fitting the error term obtained from equation 1 (ε_{it}) to the following set of equations.

$$\varepsilon_{it} = \lambda_i' F_t + e_{it} \quad (10)$$
$$F_{jt} = \rho_j F_{jt-1} + u_{jt} \quad (11)$$

$$e_{it} = \phi_i e_{it-1} + v_{it} \tag{12}$$

In equation 10, F_t is a *k* dimensional vector of common factors F_{jt} (*j*=1, …, *k*), while λ_i is a vector of factor loadings. Under the assumption of $\rho_j < 1$ for all *j*, F_t is stationary. Therefore, the integration order of ε_{it} depends only on the integratedness of e_{it}. Consequently, there is a cointegration when $\phi_i < 1$. This method allows different orders of stationarity to examine the cointegration relationship. However, the only requirement is that the dependent variable is nonstationary (Westerlund, 2008: 205). The author suggested two different test statistics as shown in the following equations.

$$DH_g = \sum_{i=1}^{n} \hat{S}_i \left(\tilde{\phi}_i - \hat{\phi}_i\right)^2 \sum_{t=2}^{T} \hat{e}_{it-1}^2 \tag{13}$$

$$DH_p = \hat{S}_n \left(\tilde{\phi}_i - \hat{\phi}_i\right)^2 \sum_{i=1}^{n} \sum_{t=2}^{T} \hat{e}_{it-1}^2 \tag{14}$$

The panel statistic (DH_p) assumes parameter homogeneity and the group statistic (DH_g) assumes parameter heterogeneity. If the calculated test statistics are above the critical value, the null of "no cointegration" will be rejected.

The CCEMG estimation method is also based on the factor modeling approach. It is an estimation method that is adding cross-section averages to the model as common factors. Pesaran (2006) suggested two estimation methods in his work. The first is CCEMG (Common Correlated Effects Mean Group), and the second one is CCEP (Common Correlated Effects Pooled), while CCEMG assumes parameter heterogeneity and CCEP parameter homogeneity. After the preliminary analysis of slope homogeneity, CCEMG estimation was found appropriate for our analysis. This method considers the cross-sectional dependency, slope heterogeneity, and non-stationarity (Eberhardt, 2012). CCEMG estimator is based on the following equation.

$$\hat{b}_i = (X_i' \dot{M_w} X_i)^{-1} X_i' \dot{M_w} Y_i \tag{15}$$

$$\hat{b}_{CCEMG} = \frac{1}{N} \sum_{i=1}^{N} \hat{b}_i \tag{16}$$

where $\hat{b}_i$ denotes individual CCE estimation for each cross-section unit. After the calculation of $\hat{b}_i$ $\hat{b}_{CCEMG}$ will be calculated by taking the average of slopes calculated for each unit. Pesaran (2006) stated that this estimator yields more effective results than the estimators that do not consider cross-sectional dependence. In addition, Chudik and Pesaran (2015) developed the dynamic version of the CCEMG estimator. This estimator is based on the idea that augmenting the CCEMG estimator with a lagged dependent variable as follows

$$y_{it} = \lambda_i y_{i,t-l} + \beta_i x_{i,t} + \sum_{l=0}^{p_T} \varphi_{i,l}' \bar{z}_{i,t-l} + \varepsilon_{it} \tag{17}$$

where $\bar{z}_{t-l}$ refers to lagged cross-sectional averages [$\bar{z}_t = (\bar{y}_t, \bar{x}_t)$]. The model with a lagged dependent variable is no more strictly exogenous. Chudik and Pesaran (2015) proposed that adding lagged cross-sectional averages mitigates the endogeneity problem. However, the authors also showed a negative bias in CCEMG estimation with a lagged dependent variable when T<50. To prevent small sample time series bias, Chudik and Pesaran (2015) suggested two different methods. The first is the split-panel jackknife method developed by Dhaene and Jochmans (2015). The second is the recursive mean adjustment method developed by So and Shin (1999). The jackknife method is based on the following equation.

$$\tilde{\pi}_{MG} = 2\tilde{\pi}_{MG} - {}^{1}\!/_{2}\,(\hat{\pi}_{MG}^{a} + \hat{\pi}_{MG}^{b}) \tag{18}$$

In equation 18, $\hat{\pi}_{MG}^{a}$ denotes the CCEMG estimation with the first half of time dimension (t = 1, 2, 3, …, (T/2)) and $\hat{\pi}_{MG}^{b}$ denotes estimation with second half of time dimension (t = (T/2)+1, (T/2)+2, …, T). Bias correction of CCEMG estimator with recursive mean adjustment method is as simple as the jackknife method. It is based on the de-meaning of the variables. According to So and Shin (1999), using the partial mean of variables is not influenced by future observations. Chudik and Pesaran (2015) revealed that the jackknife method outperforms the recursive mean adjustment method. Therefore, in our study, we used the jackknife method to prevent small sample time series bias.

Results and Discussion

The results obtained from Pesaran (2004) CD_{LM} and Pesaran et al. (2008) bias-adjusted LM (LM_{adj}) cross-sectional dependence tests and Pesaran and Yamagata (2008) slope homogeneity tests are given in table 2.

Table 2. Cross-sectional Dependence and Slope Homogeneity Results

Test	Statistics	p-value
$\mathrm{CD_{LM}}$	9.125	0.000
$\mathrm{LM_{adj}}$	59.9	0.000
$\tilde{\Delta}$	70.522	0.000
$\tilde{\Delta}_{adj}$	75.676	0.000

According to the results, the null hypotheses of cross-sectional independence are rejected at 1% significance level. The presence of cross-sectional dependency in the model indicates that there is a significant correlation between the error terms obtained for each cross-section unit. As stated earlier, using estimation methods that do not take this effect into account cause biased and inconsistent results. In addition, the results of slope homogeneity tests indicate that the null of homogeneity is rejected at 1%. This result indicates that the estimated slope parameters differ between the countries included in the analysis. In other words, the slope parameters are heterogeneous. Within the framework of this result, mean group estimators that make individual estimates for each cross-section included in the panel should be used. The CADF unit root test developed by Pesaran (2007)) is robust under cross-sectional dependency and slope heterogeneity. The results are given in Table 3.

Table 3. CADF Unit Root Test Results

	Level	1^{st} difference	Results
GDP	-2.179	-3.794***	I_1
AGR	-2.862***	-	I_0
GCF	-2.678**	-	I_0
HC	-2.054	-2.230**	I_1
TRD	-2.308	-2.752***	I_1

***, ** indicates rejection of the null hypothesis at 5% and 1% level respectively. The critical value for the model with constant and trend at 1% is -2.780 while it is -2.30 for the model with constant only. The critical value for %5 significance level for the model with constant and trend is -2.65 and for the model with constant only is -2.15.

In the unit root analysis, maximum lag is determined as 3, and optimum lag level is determined by F joint test from general to particular. While the variables are at level, the constant and trend are considered together. Only the constant is considered while the variables are first differenced. The results show that GDP, HC, and TRD variables are nonstationary at level, but their first differences are stationary. Also, AGR and GCF variables are trend stationery. These results indicate that the variables are integrated into different orders.

Our preliminary analysis shows that our variables are cross-sectionally dependent, and the slope coefficients are heterogeneous. Within the framework of these results, methods that yield robust results with variables at different orders of stationarity should be used while making cointegration analysis and long-run estimation. According to Westerlund (2008), the Durbin-Hausman cointegration method can be used while variables are integrated with different orders as long as the dependent variable is I_1. In estimating long-run coefficients, the CCEMG method proposed by Pesaran (2006) was found appropriate. The CCEMG estimator provides robust results under cross-sectional dependence, and it assumes heterogeneous slopes. It is also robust regardless of the integration levels of the variables (Eberhardt, 2012). The results of the Durbin-Hausman cointegration test are given in Table 4.

Table 4. Durbin-Hausman Cointegration Results

D-H group statistic	p-value
-1.703	0.044

In the cointegration analysis, the maximum lag level selected as 3, and optimum lag is determined by the Akaike Information Criterion. The constant term is also included in the model, and the maximum number of common factors determined as 3. Due to the results of homogeneity tests we performed, we consider the Durbin-Hausman group statistic that assumes slope heterogeneity. The results indicate that the null of no cointegration is rejected at %5 significance level. There is a significant long-run relationship between those variables.

After the cointegration analysis, the long-run coefficients are estimated via the CCEMG estimator. We also reported the Dynamic CCEMG estimation results and its bias-corrected form with a split-panel jackknife ($CCEMG_{JK}$) method. As Baltagi (2008) stated, most of the economic relations have a naturally dynamic structure, and a lagged value of the dependent variable should be among the explanatory variables in the estimation of these relations. However, a long period is required (T>50) for the Dynamic-CCEMG estimator to give unbiased results (Chudik & Pesaran, 2013). Therefore, the estimation results corrected by the half-panel Jackknife method suggested by Chudik and Pesaran (2015) are considered to be the most appropriate results. The results are given in Table 5.

Table 5. Long-Run Estimation Results

	CCEMG	*Dynamic CCEMG*	$Dynamic\ CCEMG_{JK}$
GDP (-1)		0.383 (9.83)***	0.405 (7.89)***
AGR	0.116 (3.65)***	0.082 (4.22)***	0.074 (3.11)***
GCF	0.146 (6.96)***	0.088 (4.98)***	0.097 (4.73)***
HC	0.444 (1.85)*	1.245 (1.98)**	2.515 (2.91)***
TRD	-0.023 (-1.00)	-0.009 (-0.59)	-0.003 (-0.19)
C	0.814 (1.23)	1.037 (1.20)	0.877 (0.89)

***, **, and * indicate rejection of the null hypothesis at 1%, 5%, and 10%, respectively. t statistics are given in parenthesis.

In the estimation of Dynamic CCEMG, lagged cross-sectional averages are added, and lag structure is determined as 2. According to the results of CCEMG estimation, agricultural productivity has a positive impact which is significant at 1% on GDP per capita growth. A 1% increase in agricultural productivity increases GDP per capita by 0.11%. The Dynamic CCEMG estimation and its bias-corrected version with the jackknife method are supporting the CCEMG results. A 1% increase in agricultural productivity increases GDP per capita by about 0.08%. Our results that agricultural production is one of the engines of economic growth, supporting the results of Hwa (1988), Tiffin and Irz (2006), and Yao (2000). The contribution of agricultural productivity to economic growth is strongly significant. Efforts to increase labor productivity in agriculture provide more labor input for the industrial sector while also creating an increase in demand for agricultural machinery and equipment produced by the industrial sector (Khan et al., 2020). It also increases foreign exchange earnings via rising agricultural exports. At the same time, it is necessary to increase agricultural productivity in order to feed the population employed in the industrial sector (Matsuyama, 1992). Contrary to development theories that exclude efforts to increase productivity in agriculture (Timmer, 2002) and agriculture's lower education demand compared to other sectors (Gylfason, 2000), it is seen that increasing the value of agricultural output per worker has a positive effect on the overall economic growth.

Gross capital formation is also having a positive and significant impact on economic growth at 1% level. An increase in capital formation leads to a 0.14% increase in economic growth, while dynamic estimation and bias-corrected estimation results show that its effect is 0.08% and 0.09%, respectively. Savings and investments have been seen as the main source of economic growth since the standard Neoclassical models (Solow, 1956). Human capital is also estimated as another control variable. Mankiw et al. (1992) argued that human capital should be taken into account in addition to physical capital while explaining the growth differences between countries. Supportively, our results indicate that human capital is another determinant of economic growth. CCEMG estimation shows that a 1% increase in the human capital index leads to an increase in GDP per capita by 0.44%, which is significant at 10%. The dynamic estimation results also show that the effect of human capital is positive but significant at 5% while the significance is 1% in bias-corrected estimation. Human capital formation is an important instrument for developing economies to catch up with developed economies. A high level of human capital provides a more successful and, therefore, more productive workforce in creating and implementing new technologies (Benhabib & Spiegel, 1994). According to Barro (2001), higher levels of human capital lead to lower fertility and higher investment ratios to GDP. Our results about physical and human capital formation also support the evidence of Romer (1989) and Sala-i-Martin (1997). The effect of trade openness on economic growth is the opposite of the expected. No significant relationship was found, and this result does not support the evidence of Edwards (1998), Frankel and Romer (1996), and Sachs et al. (1995) in which the authors found a significant positive impact of trade openness on economic growth.

Conclusion

In this study, the role of agriculture in economic growth is investigated with annual data of 53 middle-income countries in the period 1991-2017. As a methodological contribution to previous studies, the second-generation panel data analysis methods, which consider cross-sectional dependency, were used. First, the stationarity properties of the variables are tested with the CADF unit root test. The results show the variables are stationary in different orders. Therefore, the Durbin-Hausman cointegration test that allows different orders of stationarity was used to determine the existence of a long-run relationship. According to the Durbin-Hausman cointegration test results, there is a significant long-run relationship between economic growth, agricultural productivity, trade openness, human capital, and gross capital formation. After that, long-run coefficients are estimated via CCEMG and Dynamic CCEMG estimator. In addition, the split-panel jackknife bias correction method is used in order to deal with the small sample time series bias. The results show that agricultural growth is positively associated with overall economic growth according to CCEMG, Dynamic CCEMG, and bias-corrected CCEMG estimations. These results do not support the view that there is no need for efforts to increase productivity in agriculture as the share of agriculture in national income will naturally decrease over time. In addition to industrialization efforts, developing countries will be able to accelerate their economic growth through investments that increase productivity in agriculture. In addition to the main findings, gross capital formation and human capital also positively affect economic growth. No significant relationship was found between trade openness and economic growth.

The middle-income countries have not yet fully succeeded in their expected transformation from agriculture to industry. Naturally, there is an effort for industrialization to achieve a higher income level. However, this study shows that necessary attention should be paid to the development of the agricultural sector, just as in the industrialization to reach high income. In developing economies where agricultural production still plays an important role in economic growth, policymakers should encourage the use of technology in agriculture (Mariyono, 2009). As stated by Gardner (2005) and in the studies on production theory, the way to increase output per worker will be possible mainly by technological development and human capital investments. There is a significant difference in the context of agricultural productivity between high and middle-income countries. To cover this gap, the use of more modern methods in agriculture should be encouraged, primitive methods that are climate-dependent should be abandoned, and more resources should be allocated to agricultural research and development activities. The policymakers should provide incentives to increase infrastructure investments and transport facilities in rural areas. Besides, training opportunities should be provided to meet the need for well-educated people in agriculture to improve worker's skills. Such efforts to increase agricultural productivity may increase economic growth rates in middle-income countries. The diminishing share of agriculture in GDP and the re-allocation of workers may be good for industrial development when this re-allocation process is supported with productivity increases in the agricultural sector (Gollin et al., 2002). It is a transformation process from labor-intensive to technology-intensive production in the agricultural sector. With this transformation, countries may benefit from the positive effects of agriculture in the context of both industrial development and food production.

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Environmental Kuznets Curve: Moderating role of financial development

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Abstract

Purpose — This study analyzes the moderating role of financial development in the Environmental Kuznets Curve (EKC) hypothesis in 25 countries.

Methods — This paper uses Lin and Chu unit root test to check the stationary of the variables. The unit root test result leads to the investigation using the panel pooled mean group model.

Findings — The results of the long-run analysis show that the EKC hypothesis exists, and financial development plays its role in two ways. Firstly, it confirms the EKC hypothesis, and secondly, it improves the coefficients of supporting variables, namely economic growth, energy growth, and manufacturing value-added. The results are robust to changing the proxies of dependent as well as independent variables. The error correction model results show that the sign of the error correction term is negative and significant, implying that all of the models will converge toward their long-run equilibrium.

Implications — Financial development is a crucial determinant to reduce environmental degradation in these countries. This implies that the governments of these countries should focus on enhancing financial development for the betterment of the environment.

Originality — The study analyzes the role of the financial sector as a moderating role in the EKC hypothesis both in emerging economies and well-developed economies.

Keywords — CO2 emissions, financial development, environmental Kuznets curve, environment.

Introduction

The excellence of environment is a key subject for policymakers. Environmental degradation or greenhouse gases are professed as severe hazards especially in industrial states (Munasinghe, 1993). Ecological dilapidation also roots the progression of income insufficiency and poor sustenance in less developed economies. An improved environment delivers social firmness and economic protection to individuals (Lehtonen, 2004). Afroz, Hassan, and Ibrahim (2003) identify the adverse impact of pollution on public health.

Orru, Orru, Maasikmets, Hendrikson, and Ainsaar (2016) found that the underprivileged value of environment reduces wellbeing and pleasure of lifespan. Furthermore, the relationship

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between eminence of environment and economic performance is long lactic. This relationship is termed with a premise the environmental Kuznets curve (EKC). EKC (1955) specifies the link among numerous elements of environmental dilapidation and economic performance. In the initial phases of economic performance, greenhouse gases increase and ecological eminence decays. In the last phases, economic performance and environmental value have a positive monotonic relationship. This result infers that ecological effects or greenhouse gases are an upturned U-shaped function of economic performance indicators. The EKC is a major postulate among ecologists and environmental economists to form an assembly among ecological quality indicators and socio-economic conditions. de Bruyn, van den Bergh, and Opschoor (1998), Grossman and Krueger (1991), Panayotou (1993), and Shafik and Bandyopadhyay (1992) are among pioneer researchers who empirically tested EKC.

One of the crucial elements for growth in underdeveloped countries is the expansion of the financial sector (Patrick, 1966). The contribution of a financial segment cannot be neglected in economies (Katircioğlu & Taşpinar, 2017). An effective and well-organized financial sector supports households and businesses to apply their speculative and profit-making judgments (Shahbaz, Bhattacharya, & Mahalik, 2018). It consents economies to accomplish their insufficient means prolifically and efficiently. Financial improvement elevates business stratagems and corporate doings by offering inexpensive investment, issuing resources to dynamic segments and boosting corporations to practice the most recent tools to upturn local production.

However, the role of financial development in the context of EKC is open-ended because it may have positive and negative impacts on environmental quality. Those who reflect the positive contribution of financial development in environment dilapidation claim many arguments which are accessible here. Equally, financial development plays a vigorous part in economic evolution; it also grows power usage, which eventually causes ecological pollution (Khan, Yaseen, & Ali, 2017). On the production side, financial expansion arranges for accessibility of funding which permits organizations to purchase firsthand technology and raises energy depletion and environmental deprivation. Equally on the consumption side, financial enlargement sorts it tranquil for consumers to purchase superfluity objects like compartments and vehicles that put away energy and make a meager ecological system.

The environment Kuznets curve (EKC) was presented by Kuznets (1955). The EKC marvel is an inverted U-shape affiliation between economic growth and excellence of the environment. Kraft and Kraft (1978) are the pioneer researchers who empirically attempted to inaugurate the relation between energy consumption, economic development, and environmental depilation. Outcomes of the study reveal that economic growth is positively linked with energy consumption that persuades environmental depilation. Shafik and Bandyopadhyay (1992) empirically verified the relationship between economic growth and quality of environment. Consequently, recent studies include a mix of variables to estimate the relationship among environmental dilapidation, economic performance, and financial improvement.

Eskeland and Harrison (2003) elucidated that renowned financial division entices far-off investment in and inspires cost-effective development. Overseas companies are highly power-effectual and apply eco-friendly practices than local corporations. Well-settled financial arrangement encourages businesses to espouse up-to-date know-how in energy division, resulting in less production of energy toxins (Kumbaroğlu, Karali, & Arıkan, 2008). Capelle-Blancard and Lagun (2010) postulate that well-established financial zones strengthen commercial and corporate laws which impose certain types of penalties on companies if they do not use ecologically pleasant practices. One sort of sentence is to confine their access to advances if they produce more wastes in the troposphere. That action of financial segment raises the marketplace assessment and efficiency of the companies. Yuxiang and Chen (2011) specified a reverse association between financial expansion and industrialized impurity. They argued that financial improvement supports refining ecological eminence by growing earnings, familiarizing current knowledge, and realizing regulations concerning the safety of the environment.

Study by Tamazian, Chousa, and Vadlamannati (2009) apply feasible general least squares to consider the relationship between economic and financial progress and deterioration in ecological eminence in BRIC economies. Verdicts of the study endorse the presence of EKC. The outcomes also show that financial progress is a crucial element to drop the CO_2 per capita emissions. Sadorsky (2010) considers emerging economies to see the influence of urbanization on ecological decrepitude. The results of the study postulate a constructive role of urbanization on ecological dilapidation. Hossain (2011) found causative associations among CO_2 emissions, energy usage, economic performance, trade liberalization, and urbanization for the panel of newly industrialized countries.

Al-mulali and Binti Che Sab (2012) study a panel of thirty sub-Saharan African economies to consider the effect of energy depletion and CO_2 radiation on economic and financial growth. The fallouts reveal that energy depletion accelerates economic as well as financial growth but with the cost of high pollution. Shahbaz, Solarin, Mahmood, and Arouri (2013) found the manifestation of significant long-term interactions among CO_2 emissions, financial expansion, energy depletion, and economic progress in Malaysia. The econometric indication also specifies that financial development diminishes CO_2 emissions. Shahbaz, Khraief, Uddin, and Ozturk (2014) use ARDL approach and time series data from 1971–2010 in Tunisia to review the presence of EKC. The judgments of the study convinced long term association among economic growth, usage of power trade liberalization, and CO2 emissions. The outcomes also specified the actuality of EKC in Tunisia.

Sehrawat, Giri, and Mohapatra (2015) investigates long-run links among financial progress, economic progression, energy depletion, and ecological humiliation in India. The consequences also designate the presence of EKC for India. Salahuddin, Gow, and Ozturk (2015) examine the connection between CO_2 emissions, economic progression, financial advancement, and electricity usage in Gulf countries. The use of electricity and economic advance ought to lead to a constructive long-term affiliation with CO_2 emissions. Whereas an inverse and significant association is originated between CO_2 emissions and financial advancement. Nasreen and Anwar (2015) study dynamic panel data from 1980 to 2010 to explore the influence of financial development and energy depletion on ecological humiliation. Key judgments of the study reveal that financial improvement diminishes ecological dilapidation in high income countries and upsurges ecological dilapidation in developing economies. Hypothesis of the environmental Kuznets curve is accepted in all income panels.

Farhani and Ozturk (2015) examine long term causative association among CO_2 emissions, GDP, power utilization, financial growth, trade liberalization and urbanization in Tunisia. The fallouts of the investigation divulge an affirmative insignia for the magnitude of financial growth. The results also show two way progressive affiliations between GDP and CO_2 emissions signifying non-validity of environmental Kuznets curve (EKC) hypothesis. Dogan and Turkekul (2016) study in the USA from 1960 to 2010 to investigate the Kuznets curve (EKC) hypothesis. They found that energy depletion and urban populace upsurge ecological dilapidation. Financial growth has no consequence on it, while trade increases ecological development. In accumulation, the results do not upkeep the acceptability of the environmental Kuznets curve (EKC) premise for the USA.

A panel study of the EKC hypothesis has been done by Mironiuc and Huian (2017). They study 49 economies of Europe and Central Asia to examine the liaison among economic progress, power assimilation, financial growth, and CO_2 emanations. The study found a positive impact of energy growth, financial improvement and environmental excellences on economic progress. Saidi and Mbarek (2017) also study panel of emerging economies to investigate the effect of financial development, income, trade liberalization and urbanization on CO_2 omissions. They revealed a positive link between income and CO_2 emissions. Financial development and urbanization has negative effect on carbon emissions. Katircioğlu and Taşpinar (2017) apply second-generation method in Turkey to explore the moderating effect of financial development in a predictable environmental Kuznets curve. The contemporary study did not approve a significant regulating influence of financial growth on the effect of energy depletion on CO2 emissions in Turkey.

Study in BRICS countries by Haseeb, Xia, Danish, Baloch, and Abbas (2018) found the justification of EKC hypothesis. They also investigate that urbanization and globalization have inverse link with CO_2 emissions. Depletion of power and financial progress enhance ecological degradation. Park, Meng, and Baloch (2018) have conducted a case study of certain European Union countries to show the impact of the stimulus of internet usage, financial progress, economic progression, and trade liberalization on CO_2 emissions. They use pooled mean group estimator and panel data from 2001 to 2014. Findings show that internet usages have a positive and substantial effect on CO_2 emissions. In addition, economic progression and financial progress have a lessening deleterious effect on CO_2 emission.

Moghadam and Dehbashi (2018) explore the stimulus of financial expansion and on ecological eminence in Iran from 1970 to 2011 with annual regularities and apply ARDL approach to investigate empirical outcomes. The outcomes indicate that financial improvement speed up the dilapidation of the environs. Moreover, the outcomes did not approve with the EKC premise in Iran. ECM shows that 49% of unevenness is warranted in each pass and come up to their long-term process. Ganda (2019) uses system GMM in OECD nations to investigate the relationship between environment and financial development. The study substitutes greenhouse gases and carbon emissions and three different proxies for financial development to measure the quality of environment. The results of the study explore an adverse and substantial association between domestic credit to private sector by banks and environmental sustainability. Contrariwise, domestic credit to private sector and economic growth specifies an affirmative and momentous connection with all indicators of environmental excellences.

This current study has novelty as compared to existing studies. There are studies which study environmental Kuznets curve hypothesis (e.g. Apergis & Ozturk (2015); Dogan & Turkekul (2016); Haseeb et al. (2018); and Shahbaz et al. (2014) among others). On the other hand, there are studies which analyze the impact of financial development on environmental quality (e.g. Moghadam & Dehbashi (2018); Ganda (2019); Jalil & Feridun (2011); Sehrawat et al. (2015); Shahbaz et al. (2013); Tamazian et al. (2009); Yuxiang & Chen (2011) Tamazian et al. (2009), Yuxiang and Chen, (2010), Jalil and Feridun (2011), Shahbaz et al. (2013), Sehrawat et al. (2015), Moghadam and Dehbashi (2018), and Ganda (2019) among others). However, financial development may play an indirect role in the environmental Kuznets curve hypothesis as suggested by Katircioğlu and Taşpinar (2017) who find that financial development moderates the effect of real output on carbon dioxide. However, they discuss this role in case of Turkey only and no other countries. Their results cannot be generalized for other countries. Therefore, we are taking 25 countries which are financially emerging and developed. The present study covers this gap by checking the moderating role of financial development in EKC hypothesis by using panel data for these countries.

Methods

While dealing with panel data models, main conclusions using large cross sections and periods show that assumption of homogeneousness of slope coefficients is not rational (detail can be seen in Im, Pesaran, & Shin, 2003; M. H. Pesaran & Smith, 1995; Pesaran, Shin, & Smith, 1997, 1999). There are many methods to estimate dynamic heterogeneous panel data models with large cross sections and time periods. In the fixed effects model method, each cross section has pooled time series while intercept terms can change across cross-sectional entities. If coefficients of a slope are different, then results might not be exact. Contrarily, model can be formed on individual basis in every cross section and average values for coefficients is found. This method is called Mean Group (MG) method of estimation given by Pesaran and Smith (1995). The intercepts, error variances and coefficients of slope can vary for cross sections.

A popular technique called Pooled Mean Group (PMG) estimator developed by Pesaran et al. (1997, 1999) is used to estimate non-stationary dynamic panels. This method is based on a mixture of averaging and of amalgamating coefficients (Pesaran et al., 1997, 1999). It allows parameters of short run, error variance and intercepts to change across groups but limits the long run coefficients to be equivalent. It estimates long run coefficient as well as short run coefficients.

The general form of PMG model specification can be given as:

$$Z_{it} = \sum_{j=1}^{f} \vartheta_{ij} Z_{i,t-j} + \sum_{j=0}^{g} \partial_{ij} W_{i,t-j} + \varphi_i + \varepsilon_{it} \qquad (1)$$

Where number of cross sections is i = 1, 2, ..., n and time is t = 1, 2, 3, …, T. W_{it} is a vector of $K \times 1$ regressors, ϑ_{ij} is a scalar and φ_i represents fixed effects.
Equation (b) can be re-parameterized for getting vector error correction model as:

$$\Delta Z_{it} = \theta_i Z_{i,t-j} - \emptyset_i Z W_{i,t-j} - \sum_{j=1}^{f-1} \vartheta_{ij} Z_{i,t-j} + \sum_{j=0}^{g-1} \partial_{ij} W_{i,t-j} + \varphi_t + \varepsilon_{it} \qquad (2)$$

The error correction parameter θ_i indicates the adjustment speed. If its value is equal to zero, it means there is no long run relationship between variables. If it is negative and significant, it shows convergence of model in short run towards the long run equilibrium. The magnitude of error correction term shows the rate at which model converges.

While analyzing the models containing long time period data in dynamic panels, a very crucial problem is non-stationarity. To tackle this issue, we use Lin and Chu (LLC) and Im, Pesaran and Shin (IPS) unit root tests. Levin, Lin, and James Chu (2002) have introduced various unit root tests for panel data with different specifications depending on the assumptions about entity specific time trends and intercepts terms. The test is based on Augmented Dickey Fuller (ADF) regression for testing the issue of unit root. LLC test inflicts homogeneousness on the autoregressive coefficient (trend and intercept may change for individual series) which decides the stationarity or non-stationarity in the data. The general equation of LLC test with only intercept term is specified as:

$$\Delta z_{i,t} = \beta_{0i} + P z_{it-1} + \sum_{i=0}^{g_i} \beta_{1_i} \Delta Z_{i,t-j} + \mu_{i,t} \qquad (3)$$

In equation (c), β_{0i} is the constant term which may vary across cross sections and p stands for identical coefficient of auto-regression, β_i is order of lag, $\mu_{i,t}$ is term of disturbance assumed to be same across panels and follows ARMA process of stationarity for all cross sections.
Here:

$$\mu_{i,t} = \sum_{j=0}^{\infty} \beta \gamma_{1_i} \Delta Z_{i,t-j} + \varepsilon_{it} \qquad (4)$$

These are null and alternative hypotheses:
H_0: $\varrho_i = \varrho = 0$
H_1: $\varrho_i = \varrho < 0$ for all *i*

LLC model is based on t-statistics, where ϱ is assumed to be fixed across entities for null and alternative hypotheses.

$$t_P = \frac{\hat{p}}{SE(\hat{p})} \qquad (5)$$

Under the assumption of independence and normal distribution error term and independence of cross sections, the panel regression test statistics t_p moves to standard normal distribution with N and T $\rightarrow \infty$ and $\sqrt{(N/T)} \rightarrow 0$.

Im, Pesaran and Shin (IPS), (2003) introduced a test to check stationarity in heterogeneous panel. This test is based on ADF test for individual series. The overall test statistics is based on the average of individual series, a series can be denoted by ADF as:

$$\Delta y_{i,t} = \overline{w_i} + \theta y_{t-1} + \rho y_{it-1} + \sum_{j=1}^{p_i} p_{i,j} \Delta y_{i,t-j} + \mu_{it} \qquad (6)$$

IPS test allows heterogeneity in μ_{it} value, IPS unit root test equation is expressed as:

$$\bar{t}_T = \frac{1}{N} \sum_{i=1}^{N} t_{i,t} (p_i) \qquad (7)$$

Where $t_{i,t}$ is the ADF test statistics and pi is order of lag. In ADF, test statistics is measured as:

$$A_{\bar{t}} = \frac{\sqrt{N(T)}\,[\bar{t}_T - E(t_T)]}{\sqrt{Var(t_T)}} \qquad (8)$$

The general specification of function of CO_2 emission can be written as:

$$CO2 = f(EG, EGS, EU, MVA) \tag{9}$$

In econometric model, this can be expressed as:

$$CO2_1 = \beta_{01} + \beta_{11}EG + \beta_{21}EGS + \beta_{31}EU + \beta_{41}MVA + \mu \tag{10}$$

For robustness of above model, following models have been specified:

$$LNCO2_2 = \beta_{02} + \beta_{12}EG + \beta_{22}EGS + \beta_{32}EU + \beta_{42}MVA + \mu \tag{11}$$

$$CO2_3 = \beta_{03} + \beta_{13}EG + \beta_{23}EGS + \beta_{33}EU + \beta_{43}MVA + \mu \tag{12}$$

The model including interaction terms of financial development is as follows:

$$CO2_1 = \beta_{04} + \beta_{14}EG * FD_1 + \beta_{24}EGS * FD_1 + \beta_{34}EU * FD_1 + \beta_{44}MVA + \mu \tag{13}$$

The models used to check the robustness of model 4 by changing proxy of CO2 emissions are as follows:

$$LNCO2_2 = \beta_{05} + \beta_{15}EG * FD_1 + \beta_{25}EGS * FD_1 + \beta_{35}EU * FD_1 + \beta_{45}MVA + \mu \tag{14}$$

$$CO2_3 = \beta_{06} + \beta_{16}EG * FD_1 + \beta_{26}EGS * FD_1 + \beta_{36}EU * FD_1 + \beta_{46}MVA + \mu \tag{15}$$

The models used to check the robustness of models with interaction terms by changing the proxy of financial development are as follows:

$$CO2_1 = \beta_{07} + \beta_{17}EG * FD_2 + \beta_{27}EGS * FD_2 + \beta_{37}EU * FD_2 + \beta_{47}MVA + \mu \tag{16}$$

$$LNCO2_2 = \beta_{08} + \beta_{18}EG * FD_2 + \beta_{28}EGS * FD_2 + \beta_{38}EU * FD_2 + \beta_{48}MVA + \mu \tag{17}$$

$$CO2_3 = \beta_{09} + \beta_{19}EG * FD_2 + \beta_{29}EGS * FD_2 + \beta_{39}EU * FD_2 + \beta_{49}MVA + \mu \tag{18}$$

$$CO2_1 = \beta_{010} + \beta_{110}EG * FD_3 + \beta_{210}EGS * FD_3 + \beta_{310}EU * FD_3 + \beta_{410}MVA + \mu \tag{19}$$

$$LNCO2_2 = \beta_{011} + \beta_{111}EG * FD_3 + \beta_{211}EGS * FD_3 + \beta_{311}EU * FD_3 + \beta_{411}MVA + \mu \tag{20}$$

$$CO2_3 = \beta_{012} + \beta_{112}EG * FD_3 + \beta_{212}EGS * FD_3 + \beta_{312}EU * FD_3 + \beta_{412}MVA + \mu \tag{21}$$

We use annual panel data for twenty five courtiers over 1995 to 2017 (see Appendix). The countries have been chosen on the basis of data availability for all variables. The dependent variable is carbon dioxide (CO_2) emissions measured by three proxies, i.e. CO_2 emissions measured in metric tons per capita (CO_2(1)), CO_2 emissions measured in kilotons (CO_2(2)) and CO_2 intensity measured in kilograms of oil equivalent energy use (CO_2(3)). Independent variables are economic growth measured by GDP growth in annual percentage, squared economic growth measured by square of GDP growth in annual percentage to check the non-linear impact, energy use measured in kilograms of oil equivalent per capita (EU), manufacturing value added as a percentage of GDP (MVA) and financial development measured by three proxies, i.e. domestic credit provided by financial sector as a percentage of GDP (FD(1)), domestic credit to private sector as a percentage of GDP (FD(2)) and domestic credit to private sector by banks as a percentage of GDP (FD(3)). The variable of financial development has been used as a moderator to show the indirect link of dependent and independent variables. The data on all of the variables has been collected from World Bank's data base of world development indicators.

Results and Discussion

Table 1 presents the descriptive statistics of the variables. The value of proxy 1, proxy 2 and proxy 3 of CO_2 emission have a varies values that shown from the large range between the maximum and minimum value. It also shows that there is a constant change within the observation. Economic growth rate from for each country varies from -6.599% to 11.113%, where the negative values indicate some of countries hit decreasing in economic development. The Financial Development data, whether using proxy 1, proxy 2 or proxy 3 indicates that the values are spread out over a large range of values. This is figured from the standard deviation.

Table 1. Descriptive Statistics

	Mean	Median	Minimum	Maximum	Std. Dev.
CO_2 Emission Proxy 1	452912.8	118374.4	1800.497	5789727	1062467.
CO_2 Emission Proxy 2	8.842	8.097	2.919	24.824	4.380
CO_2 Emission Proxy 3	2.265	2.387	0.318	3.460	0.633
Economic Growth	2.471	2.449	-6.599	11.113	2.487
Energy Use	4263.999	3782.889	1052.700	18178.14	2575.566
Manufacturing Value Added	14.907	15.009	3.952	24.185	4.262
Financial Development Proxy 1	132.033	128.116	26.816	347.015	66.292
Financial Development Proxy 2	97.518	93.215	12.877	312.019	49.395
Financial Development Proxy 3	88.842	87.775	11.611	312.019	44.920

The stationarity of variables has been tested by using Im, Pesaran and Shin (IPS) and Levin, Lin and Chu (LLC) tests of panel unit root. The results found that all variables are stationary at level and others at first difference. None of the variables is stationary at second difference. As the variables are integrated at different order, therefore, it is appropriate to apply panel autoregressive distributed lag (ARDL) model for empirical analysis of the model.

Table 2. Panel ARDL (1, 1, 1, 1, 1) Long Run Results

	1	2	3	4	5	6	7	8	9	10	11	12
	Dependent Variable											
I.V.	CO21	ln CO22	CO23	CO21	ln CO22	CO23	CO21	ln CO22	CO23	CO21	ln CO22	CO23
EG	0.1920* (0.0000)	0.0102* (0.0000)	0.0554* (0.0000)									
EGS	-0.0224* (0.0000)	-0.0006 (0.2568)	-0.0065* (0.0000)									
LnEU	1.4408* (0.0012)	0.9460* (0.0000)	0.1150 (0.3903)									
MVA	-0.0515** (0.0050)	0.0002* (0.9137)	0.0030 (0.5319)	-0.0882* (0.0001)	-0.0025 (0.5764)	-0.0051 (0.2085)	0.1417* (0.0000)	0.0165* (0.0000)	-0.0007 (0.8489)	0.1473* (0.0000)	0.0168* (0.0000)	-0.0008 (0.8301)
EG* FD1				0.0012* (0.0000)	0.0002* (0.0000)	0.0002* (0.0000)						
EGS* FD1				-0.00003 (0.4367)	-0.0000 (0.2811)	-0.0000* (0.0003)						
EG* FD2							0.0011* (0.0000)	0.0001* (0.0000)	0.0003* (0.0000)			
EGS* FD2							-0.0002* (0.0000)	- 0.00003* (0.0000)	- 0.00004* (0.0002)			
EG* FD3										0.0013* (0.0000)	0.0001* (0.0000)	0.0003* (0.0000)
EGS* FD3										-0.0002* (0.0000)	-0.0000* (0.0000)	-0.0000* (0.0003)
lnEU*FD1				-0.0023* (0.0000)	-0.0002* (0.0000)	-0.0001* (0.0000)						
lnEU*FD2							-0.0014* (0.0000)	-0.0001* (0.0000)	-0.0002* (0.0000)			
lnEU*FD3										-0.0007* (0.0000)	-0.0000* (0.0000)	-0.0002* (0.0000)

* and ** show level of significance at 1% and 5%, respectively.

The long run and short run results are shown in table 2 and table 3, respectively. They have been estimated by using the technique panel autoregressive distributed lag (ARDL) proposed by Pesaran and Shin (1999) and Pesaran, Shin, and Smith (2001). We have estimated twelve models. In the first model, dependent variable is CO_2 (1), i.e. CO_2 emissions metric tons per capita and independent variables are economic growth, economic growth squared, energy use measured in kilogram of oil equivalent per capita and its natural log has been used and manufacturing value added as a percentage of GDP. The baseline purpose is to check if EKC hypothesis exists or not. The results show that it exists as the coefficient of economic growth is positive and significant and the coefficient of squared economic growth is negative and significant, showing that at earlier stages of economic growth, CO_2 emission increases but it decreases after a certain point. The findings support the results found by previous studies which are Ali, Abdullah, and Azam (2017), Dogan and Turkekul (2016), Farhani and Ozturk (2015), Javid and Sharif (2016), Katircioğlu and Taşpinar (2017), Pao and Tsai (2011), Seker, Ertugrul, and Cetin (2015), Tamazian and Bhaskara Rao (2010). They contradict with the results found by Dogan and Turkekul (2016), Farhani and Ozturk (2015), and Hong Linh and Lin (2014). The results support the findings of the sign of coefficient of energy use is positive and significant and it goes with the theory as more energy use contributes to more CO_2 emissions. However, the sign of manufacturing value added is negative and significant in specification 10 and it is against the theory but it becomes positive in the specification 11 and 12.

In equation 11 and 12 we check the robustness of results obtained about EKC hypothesis by keeping the independent variables same as in the model 10 but replace the dependent variable by two new proxies of CO_2 emissions which are CO_2 emissions measured in kilotons in equation 11 and CO_2 intensity measured per kilogram of oil equivalent energy use in model 12. The results remain the same as the EKC hypothesis still exists in both models with highly significant coefficients. The impact of energy usage is also positive and significant. One more good thing is that the sign of coefficient of manufacturing value added has become positive and significant which goes with the theory as most of the selected countries are manufacturing based economies and the manufacturing sector is a major contributor of CO_2 emissions. Although the empirical results in first three models support the theory, yet the main purpose of this study is not to check the existence of EKC hypothesis as this has been done by Dogan and Turkekul (2016), Ganda (2019), Hong Linh and Lin (2014), Tamazian and Bhaskara Rao (2010), Tang and Tan (2015).

Current study tries to test if financial development in the economy plays any role in EKC or not. This is the major gap in the literature. For this, we have taken three proxies of financial development proxied by domestic credit provided by financial sector as a percentage of GDP (denoted by FD_1), domestic credit to private sector as a percentage of GDP (denoted by FD_2) and domestic credit to private sector by banks as a percentage of GDP (FD_3). Equation 10, 11 and 12 have been estimated again using first proxy of financial development as a moderator and the results have been reported in model 13, 14 and 15. We have done the robustness analysis again as in previous three models by changing the dependent variable. The results show that the EKC hypothesis exists in one of the three specifications when financial development has been used as a moderator in the model, i.e. in equation 15 as the impact of economic growth on CO_2 emissions is positive and significant in all of the three specifications with financial development as a moderator. However, in terms of EKC hypothesis, the coefficient of squared economic growth is negative in all of the three specifications form 13, 14 and 15 but it is significant only in specification 15.

As done previously, robustness analysis for moderating impact of financial development has been conducted in specifications 16, 17 and 18 by taking a new proxy for financial development which is domestic credit to private sector as a percentage of GDP (denoted by FD_2) and taking different three dependent variables of CO2 emissions. The results show that the EKC hypothesis is validated in all of these three specifications as the coefficient of economic growth is positive while that of squared economic growth is negative in all specifications and they are highly significant. Another improvement in the results due to inclusion of financial development as a moderator is that the sign of coefficient of manufacturing value added is also positive and significant in two of the three specifications.

In the last step of robustness, we have used another proxy for financial development in specifications 10, 11 and 12. It is domestic credit to private sector by banks as a percentage of GDP (denoted by FD_3). This also confirms the existence of EKC hypothesis in develop and emerging economies as the coefficient of economic growth is positive and significant and the coefficient of squared economic growth is negative and significant. The coefficient of manufacturing value added is also positive and significant in two of the three specifications. Impact of energy use is also negative and significant in all of the specifications with financial development as a moderator. It is positive in first three specifications which are not using financial development as a moderator. This points towards another new finding that economies with better financial system can use energy in a better way to reduce its impact on CO_2 emissions.

Overall, the impact of financial development as moderator while analyzing the EKC hypothesis in emerging and well developed economies is very supportive. It is not only confirms the robust relationship using different proxies for dependent variable and financial development but also presents a realistic view of manufacturing in CO_2 emissions. Another major contribution of financial development moderator is that it improves the role of energy usage in country's CO_2 emissions. According to the best of our knowledge, there is only one study done by Katircioğlu and Taşpinar (2017) in this context for Turkey, i.e. checking moderating role of financial development in EKC hypothesis. The results found in our study support the results of this study.

Table 5. Panel ARDL (1, 1, 1, 1, 1) Short Run Results

Model	1	2	3	4	5	6	7	8	9	10	11	12
I.V.	D.V											
	CO21	ln CO22	CO23	CO21	ln CO22	CO23	CO21	ln CO22	CO23	CO21	ln CO22	CO23
ECT(–1)	-0.2666* (0.0000)	-0.4427* (0.0000)	-0.2318* (0.0000)	-0.3468* (0.0000)	-0.3162* (0.0000)	-0.2694* (0.0000)	-0.3722* (0.0000)	-0.3464* (0.0000)	-0.2692* (0.0000)	-0.3757* (0.0000)	-0.3553* (0.0000)	-0.2731* (0.0000)
DGDP	-0.0443* (0.0683)	-0.0024* (0.0619)	-0.0111* (0.0017)									
DGDPS	0.0079** (0.0413)	0.0003 (0.2191)	0.0016* (0.0112)									
DlnEU	6.8877* (0.0000)	0.5266* (0.0000)	0.3692* (0.0008)									
DMVA	-0.0657*** (0.1043)	-0.0058 (0.3435)	-0.0078 (0.3817)	-0.0334 (0.5139)	-0.0152*** (0.0667)	-0.0018 (0.8265)	-0.1089** (0.0590)	-0.0193** (0.0287)	-0.0052 (0.5717)	-0.1161** (0.0418)	-0.0188** (0.0348)	-0.0055 (0.5571)
DEG*FD1				-0.0005 (0.1359)	-0.00002 (0.1735)	-0.0000* (0.0109)						
DEG*FD2				0.0000 (0.6902)	-0.0000 (0.7073)	-0.0000** (0.0287)						
DEG*FD3							-0.0002 (0.3616)	-0.000007 (0.7744)	-0.0001* (0.0075)			
DEGS*FD1							0.00003 (0.5356)	-0.000001 (0.7009)	0.00001** (0.0345)			
DEGS*FD2										-0.0002 (0.3978)	-0.000005 (0.8387)	-0.0001* (0.0071)
DEGS*FD3										0.0000 (0.3458)	-0.0000 (0.8319)	0.0000*** (0.0662)
DlnEU*FD1				0.0008** (0.0501)	0.0001** (0.0355)	-0.00003 (0.5174)						
DlnEU*FD2							0.0020* (0.0002)	0.0002* (0.0015)	0.00009 (0.1634)			
DlnEU*FD3										0.0029* (0.0017)	0.0003* (0.0005)	0.0001** (0.0500)
C	-0.7731* (0.0001)	1.9206* (0.0000)	0.2914* (0.0000)	4.1470* (0.0000)	3.845594* (0.0000)	0.6795* (0.0000)	2.7992* (0.0000)	4.1009* (0.0000)	0.6600* (0.0000)	2.6069* (0.0000)	4.2052* (0.0000)	0.6682* (0.0000)

*, ** and *** show level of significance at 1% and 5% and 10%, respectively.

Table 3 shows the results of short run analysis. The most important discussion in short run analysis is the concept of error correction term. The results reveal that the coefficient of error correction term is negative and statistically significant in all specifications. This is an indication that model converges towards equilibrium. The results of short run analysis also show that the EKC hypothesis is confirmed in most of these specifications as the coefficient of economic growth is positive while that of squared economic growth is negative in all specifications and they are highly significant. Energy growth has positive and significant relationship with CO_2. Manufacturing value added has negative and significant relationship with CO_2. In short run, moderating role of financial development confirms the EKC hypothesis.

Conclusion

This study analyzes the moderating role of financial development in environmental Kuznets curve in 25 countries. For this, annual panel data has been collected from 1995 to 2017 by using 12 specifications. The long run relationship between variable has been used by using panel autoregressive distributed lag (ARDL) technique. The short run analysis has been found by error correction model.

Robustness of results obtained about EKC hypothesis has been done by testing different proxies; CO_2 in tons per capita, CO_2 kilotons and kilograms of oil equivalent energy use to estimate CO_2 intensity. The results remain the same as the EKC hypothesis still exists in both models with highly significant coefficients. The impact of energy usage is also positive and significant. The models have been re estimated using different proxies of financial development as a moderator and dependent variable (CO_2). The results show that the EKC hypothesis is validated in most of these specifications as the coefficient of economic growth is positive while that of squared economic growth is negative in all specifications and they are highly significant. Another improvement in the results due to inclusion of financial development as a moderator is that the sign of coefficient of manufacturing value added is also positive and significant in two of the three specifications.

The coefficient of manufacturing value added is also positive and significant in two of the three specifications. Impact of energy use is also negative and significant in all of the specifications with financial development as a moderator. It is positive in first three specifications which are not using financial development as a moderator. This points towards another new finding that economies with better financial system can use energy in a better way to reduce its impact on CO_2 emissions.

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Appendix

List of the Countries

1. Australia	14. Luxembourg
2. Austria	15. Mexico
3. Belgium	16. Netherlands
4. Chile	17. Norway
5. Czech Republic	18. Poland
6. Denmark	19. Portugal
7. France	20. Spain
8. Germany	21. Sweden
9. Hungary	22. Switzerland
10. Iceland	23. Turkey
11. Israel	24. United Kingdom
12. Italy	25. United States.
13. Japan	

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Drivers of business cycles in Iran and some selected oil producing countries

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Abstract

Purpose — This study is aimed at analyzing the main drivers of business cycle in Iran and some selected oil producing countries during the 1970:Q1-2015:Q4 period. In addition, the study evaluates causality of leading macroeconomic indicators for each different regimes of the business cycles.

Methods — This study proposes a new methodological approach by combining Markov-Switching Vector Autoregressive (MSVAR) and MS-Granger causality approach.

Findings — The results show that there are diverse sources of business cycle. Iran experienced higher volatility of GDP where machinery investment and export are found as main driver of its business cycle. Meanwhile, consumer price index has countercyclical effect in all countries. We also find some similarities to the US, the UK, and Canada regarding the probability of a business cycle, number of observations, and the average duration, especially in the first regime of MS-VAR models. The high level of oil price volatility relative to the GDP volatility indicates the power of oil price shock to generate cycles. In addition, the results of the traditional Granger causality test confirm the Markov-Switching Granger Causality (MS-GC) test in all countries except export from the UK.

Implication — Identification the main driver of business cycles is very significant to formulate the steady growth path so that the government able to select the most adequate economic policy.

Originality — The novelty of this study is the adoption of a new approach by combining stylized facts and MS-VAR and MS-Granger causality to analyze the business cycles in different regime.

Keywords — Business cycles, causal variables, MSVAR, MS-Granger Causality.

Introduction

Business cycles are defined as frequent and broad-based movements in an aggregate economic activity where expansionary periods are pursued by contraction (Burns & Mitchell, 1946). The empirical relationships of business cycles between output and other economic variables are often referred to as "stylized facts" (Kydland & Prescott, 1990). As mentioned by Alp and Kilinç (2014), it is necessary for policymakers to understand the sources and properties of business cycles and to

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develop more structural models. For example, documentation of the stylized facts of business cycles is an important tool for constructing theoretical models as statistical benchmarks and for evaluating the validity of the different theoretical models.

According to Diebold and Rudebusch (1996) the Burns and Mitchell's definition of business cycles has two key features. The co-movement between individual economic variables is the first feature. Indeed, the co-movement among series was the centerpiece of Burns and Mitchell's methodology, considering possible leads and lags in timing. Burns and Mitchell's second significant factor in defining business cycles is their division of business cycles into separate phases or regimes. Many of existing studies have witnessed a synthesis of co-movement and nonlinearity features of cycles, such as Diebold and Rudebusch (1996), Chauvet (1998), Carriero and Marcellino (2007), and Buss (2010). There is room for analysis by incorporating both factor structure and regime-switching. In fact, these studies indicated that it could be useful in evaluating the turning points in business cycles if the leading indicators were to be applied to Markov-switching models (MS).

In the same context, Moradi (2016) argues that the formation of dynamic factor models and the composition of indices resulted in the first feature and the second one inspired the use of nonlinear regime-switching models with the seminal work of Hamilton (1989). Since Hamilton's model of the US business cycle until recent years, the Markov-switching autoregressive model has become increasingly popular for the empirical characterization of macroeconomic fluctuations. Various studies examining the business cycle dynamics of economic growth to date have estimated the regime-switching models for example Medhioub and Eleuch (2013), Burzala (2012) and Billio, Ferrara, Guégan, and Mazzi (2013) for the surveys on the regime-switching models. These studies are based on the movement of single series such as Real Gross Domestic Product (GDP) or Industrial Production Index (IPI).

For considering co-movement and estimating a common regime probability of a set of variables, Krolzig (1997) proposes a multivariate extension of this model by developing an MS-VAR process. Clements and Krolzig (2003) addressed the characterization and testing of business cycle asymmetries based on models of MS-VAR in later studies. By using a multi-move Gibbs sampler, (Pelagatti, 2011) estimated a duration-dependent MS-VAR model because the computational burden of using the MLE approach for such models is high.

Lee, Liang, and Chou (2013) argue that the MS-VAR model carries the characteristics of the univariate MS model with its capabilities to differentiate cycle states and capture the sustainability of states. It also reflects the co-movement between the time series and economy series in which prior univariate MS models failed to address. However, since the direction of causality between the variables is not clear, co-movements may only serve as a starting point for further studies purposing to acknowledge the causal relationships. The MS-VAR model is very useful to spot the causal relationships and treat the changes in causality as random events governed by the Markov process such that it could capture the instability of Granger causality between variables. It is a kind of VAR model where the intercept, parameter coefficients, and error term are all subject to Markov-switching.

Understanding the causes of the cyclical fluctuations is very significant, since better economic planning subjects to recognition of the causes of this fluctuation. On the other hand, understanding the causes of the cyclical fluctuations helps to find a steady growth path. Therefore it is important to analyze the causes or driving forces behind the economic fluctuations to select the most adequate government policy.

According to the empirical study by Camacho and Perez-Quiros (2002), the combination of MS models and non-parametric models derived the best out-of-sample forecasting performance. The objective of this paper is to determine the stylized facts of business cycles of countries namely Iran, the US, the UK, and Canada, within the period 1970:Q1-2015:Q4. The paper also determines the foremost cyclical characteristics like volatility and co-movement to identify the predominant drivers of business cycles using the most important leading indicator series. Furthermore the paper identifies the causal relationship between the variables using the MS-VAR method to enhance the understanding of the business cycle.

Methods

The quarterly data sources from the websites of the Federal Reserve Bank of St. Louis, World Bank Open Data, the Main Economic Indicators published by the OECD, and the Central Bank of Iran. The central bank of Iran have elaborated quarterly series of the component of aggregate demand and supply since 1988. Quarterly data between 1970 and 1988 was obtained through the method of a related series of Chow and Lin (1971). The data are adjusted seasonally since we are interested in the percentage (rather than absolute) deviations from trend. All data are expressed in logarithms and a cyclical component of variance is obtained through the double Hodrick and Prescott (1997) filter.

In this paper, we are going to use the double-HP approach suggested by the OECD system of leading indicators. Following the framework described in Arby (2001), Arby (2001), the HP filter is employed in two steps to separate these components. Firstly, the time series is decomposed and long-run trend (T_t) is eliminated by subtracting the trend from the original series (y_t). We get a new series (Z_t) that contains the cyclical and irregular component:

$$Z_t = y_t - T_t = C_t + I_t \tag{1}$$

In the second step, with using again the HP filter on $(\mathrm{Z_t})$, we will obtain the smooth component which is a cycle (C_t). The difference between $(\mathrm{Z_t})$ and (C_t) demonstrates shocks or irregular component $(\mathrm{I_t})$ calculated conforming to the following equation:

$$Min \{\textstyle\sum_{t=1}^{T}(Z_t - C_t)^2 + \lambda \sum_{t=2}^{T-1}[(C_{t+1} - C_t) - (C_t - C_{t-1})]^2\} \tag{2}$$

$$I_t = Z_t - C_t \tag{3}$$

Selection of Leading Variables

In this section, we move on to derive the leading variables among 15 macroeconomic variables for each country. In particular, the focus is on two main statistics considered in the literature as standard statistics to elucidate business cycle attributes of the related times series: (1) Relative volatility, defined as the standard deviation of each variable relative to the standard deviation cyclical component of GDPs (δ_x/δ_y). The relative volatility coefficient measures how unstable is variable x with respect to GDP. It corresponds to the ratio between the standard deviation of the x cycle and the standard deviation of the GDP cycle. As stated by Kamil and Lorenzo (2005) levels of cyclical volatility are classified according to the following convention: high (relative volatility greater than 2), medium (relative volatility greater between 1 and 2), and low (relative volatility smaller than 1). A highly volatile variable cannot be associated immediately with causality unless it also presents a cyclical pattern. (2) Co-movement, defined as the degree of contemporaneous co-movement of the variable relative to GDP. Co-movement analysis is typically made up of two aspects: time and direction. In terms of time, phase change of the variable with respect to the reference cycle series could also be leading, coinciding or lagging series. Referring to direction, they will be procyclical, countercyclical or a-cyclical. Following Agénor, McDermott, & Prasad (2000), the degree of co-movement of a series x_t with GDP (y_t) is by the magnitude of the coefficient of correlation $\rho(j)$, j∈{0,±1,±2...}. It is derived from our series using the same filter HP. The series x_t considered to be procyclic, acyclic or countercyclical if the contemporaneous correlation $\rho(0)$ is positive, zero or negative, respectively. If some of the significant cross-correlation coefficients are negative and some are positive and the largest of these are close to each other, then the cyclical properties of this variable are not clear. Furthermore, if $|\rho(j)|$ is a maximum for a positive "*j*" we say that x_t leads the cycle by "*j*" periods, is coincident if $|\rho(j)|$ is a maximum for j=0, and lags the cycle if $|\rho(j)|$ is a maximum for a negative "*j*". If the co-movement pattern of the variables is not clear then the phase shift will be also not clear. Furthermore, if the counter cyclicality or procyclicality is observed both at positive and negative periods at similar levels, then the phase shift is going to be again not clear.

Markov-Switching Vector Auto Regression (MS-VAR)

In the Markov-Switching Vector Autoregressive Model (MS-VAR), the variables under examination change their behaviors during the time, i.e. switches between regimes. The basic idea behind this class of regime-switching models is that the parameters of the VAR process will be regime dependent or depend on the state (regime) variable (s_t) which is an unobservable variable. In other words, the parameters of a K-dimensional vector time series process $\{y_t\}$ depend on an unobservable regime variable $s_t \in \{1, \ldots, M\}$, which represents the probability of being in a particular state of the world.

$$p(y_t|Y_{t-1}, s_t) = \begin{cases} f(y_t|Y_{t-1}, \theta_1) & if\ s_1 = 1 \\ \vdots & \vdots \\ f(y_t|Y_{t-1}, \theta_M) & if\ s_t = M \end{cases} \tag{4}$$

Where Y_{t-1} denotes all the observations of $[\,y_{t-j}]_{j=1}^{\infty}$ $and\ \theta$ show the parameters of the VAR model. In every regime y_t is generated by a VAR process of order q as follows:

$$y_t = \mu(s_t) + \sum_{i=0}^{q} A_i(s_t) y_{t-i} + u_t \tag{5}$$

$$u_t \sim NID(0, \sum(s_t))$$

Where $\mu(.)$ shows the intercepts or mean in each regime, $A_i(.)$ is a matrix and shows the coefficients of the lagged values of the variable in different regimes, and $\sum$ shows the variance of the residuals in each regime. To complete the data generating process assumptions about the regime generating process (s_t) are needed.
In MS-VAR models the regime-generating process (s_t) is produced by a Markov chain:

$$Pr[s_t|\{s_{t-1}\}_{i=1}^{\infty}, \{y_{t-1}\}_{i=1}^{\infty}] = Pr\{s_t|s_{t-1}; \rho\} \tag{6}$$

Where ρ includes the probability parameters. The result is that the current regime (s_t) depends only on the regime one period ago (s_{t-1}). Thus the transition probabilities could be shown as:

$$P_{ij} = P\left\{s_{t=j}\middle|s_{t-1} = i, s_{t-2} = k, \ldots\right\} = \Pr\left\{s_{t=j}\middle|s_{t-1} = i\right\} \tag{7}$$

$$, \sum_{j=1}^{M} P_{ij} = 1\ \forall i, j \in \{1, \ldots, M\}$$

Where (P_{ij}) gives the probability that state (*i*) will be followed by state (*j*) and $0 \leq P_{ij} \leq 1$. These transition probabilities can be indicated in an $(M \times M)$ transition matrix:

$$p = \begin{bmatrix} p11 & p12 & \cdots & p1M \\ p21 & p22 & \cdots & p2M \\ \vdots & \vdots & \vdots & \vdots \\ pM1 & pM2 & \ldots & pMM \end{bmatrix} \tag{8}$$

According to Krolzig (1997), in the general specification of MS-VAR model, all autoregression parameters are conditioned on the state (s_t) of the Markov chain. So, equation (9) can be written as:

$$y_t = v(\varepsilon_t) + A_1(\varepsilon_t)y_{t-1} + \cdots + A_p(s_t)y_{t-p} + u_t \tag{9}$$

Pursuant to Krolzig (1997) in the most general specification of an MS-VAR model, all parameters of the auto regression are conditioned on the state (s_t) of the Markov chain such that each regime $m\ VAR(p)$ parameterization $v(m)$ (orμ_m), $\Sigma_m, A_{1m}, \ldots, A_{jm}, m = 1, \ldots, M$ such that:

$$y_t = \begin{cases} v_1 + A_{11}y_{t-1} + \cdots + A_{p1}y_{t-p} + \sum_1^{1/2} u_t\ if\ s_1 = 1 \\ \vdots \\ v_M + A_{1M}y_{t-1} + \cdots + A_{pM}y_{t-p} + \sum_M^{1/2} u_t\ if\ s_1 = M \end{cases} \tag{10}$$

Where $u_t \sim NID(0, I_k)$.

Estimation of MS-VAR models are in many empirical analyses based on the EM algorithm suggested by Hamilton (1989). The EM algorithm has been designed to estimate the parameters of a model where the observed time series depends on an unobserved or a hidden stochastic variable. The iterative estimation technique can be used to make inference for $t = 1,2, \dots, T$, while taking the previous value of this probability $\xi_{it-1} = P_r\{s_{t-1} = i|\Omega_{t-1}; \theta\}$ as an input. The conditional log-likelihood can be given by:

$$logf(y_1, y_2, \dots, y_T|y_0; \theta) = \sum_{t=1}^{T} logf(y_t|\Omega_{t-1}; \theta) \quad (11)$$

Markov Switching Granger causality (MS- GC)

The methodology requires the estimation of either an MSIA(.)-VAR (.) or an MSIAH (*p*)-VAR (*q*) model. Based on the coefficients of the lagged values we can determine the direction of the Granger causality in the equation for each variable. For example, in countries with three variables the model is given as:

$$\begin{bmatrix} y_{1t} \\ y_{2t} \\ y_{3t} \end{bmatrix} = \begin{bmatrix} \mu_{1,s_t} \\ \mu_{2,s_t} \\ \mu_{3,s_t} \end{bmatrix} + \sum_{k=1}^{p} \begin{bmatrix} A_{11,s_t}^{(k)} & A_{12,s_t}^{(k)} & A_{13,s_t}^{(k)} \\ A_{21,s_t}^{(k)} & A_{22,s_t}^{(k)} & A_{23,s_t}^{(k)} \\ A_{31,s_t}^{(k)} & A_{32,s_t}^{(k)} & A_{33,s_t}^{(k)} \end{bmatrix} \begin{bmatrix} y_{1,t-k} \\ y_{2,t-k} \\ y_{3,t-k} \end{bmatrix} + \begin{bmatrix} \epsilon_{1,s_t} \\ \epsilon_{2,s_t} \\ \epsilon_{3,s_t} \end{bmatrix} \quad (12)$$

In the y_{1t} vector, y_{2t} and/or y_{3t} is/are Granger cause of y_{1t} in each k^{th} regime if the parameter set or sets of $A_{12}^{(k)}$ and $A_{21}^{(k)}$, and/or $A_{13}^{(k)}$ and $A_{31}^{(k)}$ are statistically different from zero. In general, Granger causalities can be detected by testing $H_0: A_{12}^{(k)} = 0$ and $H_0: A_{21}^{(k)} = 0, H_0: A_{23}^{(k)} = 0$ and $H_0: A_{32}^{(k)} = 0,$ and $H_0: A_{13}^{(k)} = 0$ and $H_0: A_{31}^{(k)} = 0.$

Results and Discussion

Selection of Causal Variables

As mentioned by Pacheco Jiménez (2001) the stylized facts is very beneficial to evaluate causal relationships because a typical "causal" variable shows a leading, procyclical, and highly volatile behavior. This suggests, but not confirms, causality respect to GDP. Table 1 provides the statistical relationships between real GDP and 15 other variables. The variables are seasonally adjusted and taken logarithms before calculating the volatility and correlation.

The results indicate that the volatility of GDP in Iran is much higher than other industrialized countries studied in the study. Based on Male (2011), higher output volatility reflects, the vulnerability of Iran's developing economy and its inability to diversify risks or perform stabilizing macroeconomic. Moreover, consumption, investment, exports and imports in all countries are found to be often procyclical. On the other hand, the consumption of durable goods and machinery investment has a high relative volatility. Government final consumption in Iran, unlike other countries, has a positive contemporaneous correlation and also procyclical.

The relative volatility in the unemployment rate is high in all countries, but its contemporaneous correlation is negative and it does not show a clear pattern. Regarding monetary variables, it can be mentioned that narrow money in the UK and Canada are recognized as the main driver of the business cycle. The consumer price index (CPI) has low relative volatility, negative contemporaneous correlation, and countercyclical behavior in all countries. The high level of oil price volatility relative to the GDP volatility in all countries, except for Iran, implies the ability of oil price shock to generate cycles. Meanwhile the lower level in Iran, perhaps can be explained simply by the extremely high output volatility experienced in Iran. On the other hand, the phase shifts in all countries are coincident or not clear.

In the next step, the integration order of the variables were determined using the point optimal test of Hylleberg, Engle, Granger, and Yoo(1990)–HEGY test and Ng and Perron (2001)–NGP test were determined before testing causality. In the second step, the maximum likelihood procedure of Johansen is utilized for the determination of the possible existence of cointegration between variables.

Table 1. Statistical relationships between real GDP and other variables

	US	Canada	UK	Iran
Real GDP				
Volatility	1.428	1.332	1.505	7.289
Autocorrelation (t, t-1)	0.942	0.928	0.939	0.954
Private consumption expenditure				
Relative Volatility	0.836	0.522	0.952	0.680
Contemporaneous Correlation	0.904	0.761	0.032	0.562
Cyclicality	Procyclical	Procyclical	countercyclical	Procyclical
Phase Shift	Lead (.921)	Coincidental	lag(-0.241)	lag(-0.577)
Durable goods				
Relative Volatility	2.402	2.368	2.634	2.957
Contemporaneous Correlation	0.814	0.740	0.672	0.108
Cyclicality	Procyclical	Procyclical	Procyclical	countercyclical
Phase Shift	Lead (0.876)	Lead (0.748)	Coincidental	Lag (0.259)
Non-durable goods				
Relative Volatility	1.255	0.501	1.235	0.660
Contemporaneous Correlation	0.270	0.698	0.464	0.469
Cyclicality	Procyclical	Procyclical	Procyclical	Procyclical
Phase Shift	lag(.445)	Lag (0.733)	Lag (0.481)	Lag (0.530)
Services				
Relative Volatility	0.503	0.690	0.820	0.945
Contemporaneous Correlation	0.335	0.786	0.888	0.442
Cyclicality	Procyclical	Procyclical	Procyclical	Procyclical
Phase Shift	Coincidental	Coincidental	Coincidental	Coincidental
Fixed Investment				
Relative Volatility	2.972	2.642	2.759	1.547
Contemporaneous Correlation	0.962	0.732	0.873	0.436
Cyclicality	Procyclical	Procyclical	Procyclical	Procyclical
Phase Shift	Coincidental	Coincidental	Coincidental	Lead (0.442)
Construction Investment				
Relative Volatility	6.770	2.329	3.614	1.362
Contemporaneous Correlation	0.725	0.652	0.633	0.232
Cyclicality	Procyclical	Procyclical	Procyclical	Procyclical
Phase Shift	Lead (.852)	Coincidental	Coincidental	Lead(0.493)
Machinery investment				
Relative Volatility	3.669	4.427	3.635	2.996
Contemporaneous Correlation	0.841	0.706	0.637	0.232
Cyclicality	Procyclical	Procyclical	Procyclical	Procyclical
Phase Shift	Lead (.866)	Coincidental	Lag (.654)	Lead (.310)
Government final consumption				
Relative Volatility	0.685	0.656	0.500	0.724
Contemporaneous Correlation	-0.366	-0.222	-0.289	0.531
Cyclicality	countercyclical	countercyclical	countercyclical	Procyclical
Phase Shift	Coincidental	Lag (-0.363)	Lag (-0.312)	Lag (.546)
Exports				
Relative Volatility	2.731	2.741	2.050	2.336
Contemporaneous Correlation	0.498	0.767	0.543	0.566
Cyclicality	Procyclical	Procyclical	Procyclical	Procyclical
Phase Shift	Lag (-0.634)	Lag (.768)	Lead (.558)	Lead (.611)
Imports				
Relative Volatility	3.284	3.184	2.188	1.893
Contemporaneous Correlation	0.878	0.813	0.800	0.417
Cyclicality	Procyclical	Procyclical	Procyclical	Procyclical
Phase Shift	Coincidental	Coincidental	Coincidental	Coincidental
Narrow Money (M1)				
Relative Volatility	6.488	2.133	2.018	0.893
Contemporaneous Correlation	0.009	0.249	0.598	0.207
Cyclicality	Acyclical	Procyclical	Procyclical	Procyclical
Phase Shift	Not Clear	Lead (0.343)	Lead (0.679)	Coincidental
Broad Money (M3)				
Relative Volatility	0.854	1.545	2.036	0.800
Contemporaneous Correlation	0.101	0.147	-0.118	0.208
Cyclicality	Procyclical	Not Clear	countercyclical	Procyclical
Phase Shift	Lead (0.347)	Not Clear	Lead (-0.476)	Lead(0.213)
CPI				
Relative Volatility	0.908	0.977	1.144	1.369
Contemporaneous Correlation	-0.458	-0.513	-0.709	-0.210
Cyclicality	countercyclical	countercyclical	countercyclical	countercyclical
Phase Shift	Lead (-0.689)	Lag (-0.638)	Coincidental	Lead (-0.328)
Oil Price				
Relative Volatility	11.236	12.396	10.898	3.099
Contemporaneous Correlation	0.116	0.322	0.107	0.338
Cyclicality	Not Clear	Procyclical	Not Clear	Procyclical
Phase Shift	Not Clear	Coincidental	Not Clear	Coincidental
Unemployment				
Relative Volatility	8.077	6.254	5.391	9.569
Contemporaneous Correlation	-0.894	-0.851	-0.722	-0.060
Cyclicality	countercyclical	countercyclical	countercyclical	Acyclical
Phase Shift	Lead (-0.918)	Lag (-0.865)	Lag (-0.823)	Not Clear

The results indicate that the null hypothesis of a unit root cannot be rejected at the 5 % level of significance. On the other hand, the first differences of variables appear to be stationary. As a result, we can say that the variables are integrated of order one, I(1). Since the variables are integrated of the same order, the maximum likelihood procedure of Johansen can be used to examine the possible existence of cointegration between the variables. According to the results, the null hypothesis of no cointegration could not be rejected. Because they are not cointegrated, the first difference or innovations of the variables can be used to test for MS- Granger.

MS-VAR and MS-Granger Causality Test Results

The first difference or innovations of the variables is used to Markov Switching- Granger causality analysis. Before estimating the MS models, a linear VAR model is specified. The lag length is determined by using information criteria such as AIC and SIC. Next, the MSIA and MSIAH models are estimated for each country using the selected optimal lags assuming two and three regimes. In these models, for determining the number of regimes, we firstly tested linear VAR model against an MS-VAR model with two regimes. The H_0 hypothesis will reject linear hypothesis by using LR test statistics for all countries.

Next, MS-VAR model with two regimes are tested against its alternative with three regimes. The null hypothesis, which indicates superiority of model with two regimes, was rejected based on calculated LR statistic.

Iran

The MSIA(2)-VAR(4) model, which is the first model to be analyzed, is estimated for the IRAN and the results are given in Table 2. Based on the transition probabilities, the most persistent regime is regime number 2 and the probability of staying in this regime is 67%. On the other hand, the tendency to remain in regime 1 is extremely low (14%). The business cycle phases with the majority of observations (129 quarters) exist in regime 2.

Table 2. Regime properties of the MSIA (2)-VAR (4) model for IRAN

	Number of Observations	Probability	Average Duration	Transition probability		
Regime 1	43	25.00	1.13	Regime 1:	0.14	0.86
Regime 2	129	75.00	3.31	Regime 2:	0.33	0.67

To find the potential similarities and differences of causality, these results are compared with traditional causality tests in Table 3.

Table 3. MS-Granger and Linear Granger causality results for IRAN

Regime 1			Regime 2		
Exp	↔	gdp	Exp	←	gdp
Mach	↔	gdp	Mach	none	gdp

Causality Direction	χ^2	Prob.	Causality Decision
Δgdp → Δexp	3.70	0.44	No
Δgdp → $\Delta mach$	5.69	0.22	No
Δexp → Δgdp	71.39	0.00	Yes
$\Delta mach$ → Δgdp	8.43	0.07	Yes

The results indicate that there is unidirectional traditional Granger causality from export to GDP, while there is bidirectional MS-Granger causality between exports to GDP in regime 1 and unidirectional from GDP to exports in regime 2. Furthermore, there is evidence to support unidirectional traditional Granger causality from machinery investment to GDP, while there is bidirectional MS-Granger causality between machinery investment and GDP in regime 1.

United States

The estimation results of MSIAH(2) -VAR(3) model for the US are given in Table 4. As it's clear, probability of remaining in regime 1 is calculated at 66 % while the probability of shifting to regime 2 is 34%. The possibility of proceeding to the first regime from the second regime is very high (76%), whereas the possibility of staying in the second regime is 24%.

For comparative purposes, the traditional linear Granger causality test results and summary of the MS-Granger causality test results are exhibited in Table 5.

Table 4. Regime properties of the MSIAH (2)-VAR (3) model for the US

	Number of observation	Probability	Average Duration	Transition probability
Regime 1	120	69.77	3.00	Regime1: 0.66 0.34
Regime 2	52	30.23	1.30	Regime2: 0.76 0.24

Table 5. MS-Granger and Linear Granger causality results for the US

Regime 1		Regime 2	
dur ↔ gdp		dur ↔ gdp	
con → gdp		con ← gdp	
mach → gdp		mach ← gdp	
Causality Direction	χ^2	Prob.	Causality Decision
Δgdp → Δdur	3.99	0.26	No
Δgdp → Δcon	3.32	0.34	No
Δgdp → Δmach	7.00	0.07	Yes
Δdur → Δgdp	10.93	0.01	Yes
Δcon → Δgdp	27.09	0.00	Yes
Δmach → Δgdp	20.68	0.00	Yes

Based on the traditional Granger causality test, there is a unidirectional relationship running from durable consumption to GDP, but the bidirectional relationship is accepted in all regimes. Whereas for the causal relationship between GDP and construction investment, we found traditional Granger causality running from construction investment to GDP. On the contrary unidirectional relationship from GDP to construction investment is only found in regime 2. Moreover, there is a bidirectional traditional Granger causality between GDP and machinery investment. However the results of MS-Granger causality tests show that there is a unidirectional relationship running from machinery investment to GDP in regime 1 and from GDP to machinery investment in regime 2.

Canada

Table 6 portrays the MSIAH(3)-VAR(2) model as the best fit to Canadian data. In this country, the business cycle phase with the most duration is the first phase (19 quarters on average). The transition probabilities suggest the persistence of regime (1) is higher than other ones (0.85).

Table 6. Regime properties of the MSIAH (3)-VAR (2) model for CANADA

	Number of Observations	Probability	Average Duration	Transition Probability
Regime 1	95	69.34	19.00	Regime1: 0.95 0.04 0.01
Regime 2	21	15.33	3.00	Regime2: 0.24 0.68 0.08
Regime 3	21	15.33	7.00	Regime3: 0.00 0.15 0.85

In order to find the potential similarities and differences of causality, these results are compared with traditional causality tests in Table 7. The results indicate that there is a bidirectional traditional Granger causality between GDP and durable consumption, which corresponds with the results of MS-Granger causality tests in regime 1. Moreover, there is unidirectional traditional

Granger causality from narrow money to GDP which is consistent with results obtained for regimes 1 and 3.

Table 7. MS-Granger and Linear Granger causality results for the CANADA

regime 1	regime 2	regime 3
dur ↔ gdp	dur None gdp	dur None gdp
m_1 → gdp	m_1 ← gdp	m_1 → gdp

Causality Direction	χ^2	Prob.	Causality Decision
Δgdp → Δdur	6.64	0.03	Yes
Δgdp → Δm_1	2.98	0.22	No
Δdur → Δgdp	12.52	0.00	Yes
Δdur → Δm_1	2.76	0.25	No
Δm_1 → Δgdp	5.89	0.05	Yes
Δm_1 → Δdur	17.08	0.00	Yes

United Kingdom

Table 8 represents the results of estimated MSIAH(3)-VAR(3) model for UK. The transition probability matrix shows that the first two regimes are more persistent than regime 3. Furthermore, regimes 1 and 2 include 89 and 75 quarters, respectively.

Table 8. Regime properties of the MSIAH (3)-VAR (3) model for the UK

	Numberof observation	Probability	Average Duration	Transition probability
Regime 1	89	49.44	4.24	Regime1: 0.75 0.18 0.07
Regime 2	75	41.67	3.95	Regime2: 0.24 0.75 0.01
Regime 3	16	8.89	2.67	Regime3: 0.12 0.26 0.62

The comparison between traditional linear Granger causality test and the MS-Granger causality test are exhibited in Table 9.

Table 9. MS-Granger and Linear Granger causality results for the UK

regime 1	regime 2	regime 3
exp None gdp	exp ↔ gdp	exp ↔ gdp
m_1 ↔ gdp	m_1 ↔ gdp	m_1 → gdp

Causality Direction	χ^2	Prob.	Causality Decision
Δgdp → Δexp	**12.25**	**0.00**	**Yes**
Δgdp → Δm_1	**14.09**	**0.00**	**Yes**
Δexp → Δgdp	0.37	0.94	No
Δm_1 → Δgdp	**22.09**	**0.00**	**Yes**

In line with the traditional Granger causality test, there is a unidirectional relationship running from GDP to export. The bidirectional relationship in the MS-Granger causality test exhibit in regime 2 and regime 3. Moreover, bidirectional traditional Granger causality between GDP and narrow money is found for regimes 1 and 2 while, our findings confirmed existence of unidirectional causality from narrow money to GDP in regime 3.

Key Findings of Estimated MS Models

The probability of the business cycle remaining in regime 1 was longer than other regimes. The business cycle phases with the number of observations of each regime's existence in regime 1 suggest the persistence of this regime in the US, Canada, and the UK. The average duration shows a similar pattern in the business cycle in all countries except Iran. The longest average duration with 19 quarters is registered in the first regime of Canada and the highest number of observation with 129 quarters in regime 2 is recorded in Iran. The transition probability matrix shows that all the regimes are persistent in all countries except the probability of remaining in the second regime in the US and the possibility of staying in the first regime of Iran.

In general business cycle in Iran is relatively different to the other OECD countries. The factors underlying this difference are an Iran's single-product (oil) economy, different degrees of development and the lack of a strong business relationship with foreign countries, especially after the revolution in Iran. As a result, these factors lead to restrain in the exports of oil, imports of investment in intermediate goods, extra financial costs, and severe fluctuations in Iran's GDP.

According to the results obtained by the MS-VAR and MS-Granger causality approach, all studied variables are recognized as the main driver of business cycle at least in one regime. Nevertheless, the results of the traditional Granger causality test confirm the MS-Granger causality test results in all countries except export from the UK.

The result of the first step in trying to identify the main driver of the business cycle using some selected macroeconomic time series is similar to some previous works including De Medeiros and Sobral (2011), Camacho and Perez-Quiros (2002), and Misas and Ramírez (2007). Findings of the second step of recognizing the direction of causality are in lin e to those studies based on MS-VAR and MS-Granger causality test like Bildirici (2013), Hyera and Mutasa (2016), Claessens, Kose, & Terrones (2012) and Billio, Anas, Ferrara, & Duca (2007). No prior studies, however, have highlighted both steps and the novelty of this paper lies in the combination of both steps.

Conclusion

This study attempted at uncovering drivers of business cycles by presenting evidence from Iran, the US, the UK, and Canada for the 1970:Q1-2015:Q4 period. The first objective is to find the main driver causing the business cycle by using the stylized facts of business cycles of countries surveyed. The second goal of the study is to recognize the direction of causality, based on MS-VAR and MS-Granger causality approach to evaluate causality in different regimes of the business cycle and compares the result with the traditional Granger causality test.

The results revealed higher volatility of GDP in Iran compared to three developed countries. However, we find some similarities within the US, the UK, and Canada regarding the probability of a business cycle, number of observations, and the average duration, especially in the first regime of MS-VAR models. Export and machinery investment in Iran are Granger causes of GDP in the only regime 1. The causal variable in the US could be durable consumption, construction investment, and machinery investment. MS-Granger analysis results show that durable consumption is the Granger cause of GDP in all regimes, but machinery investment and construction investment are Granger causes of GDP in the only regime 1. According to the results obtained for Canada, durable consumption and narrow money can be considered as Granger causes of GDP in regime 1. Meanwhile, export and narrow money play the same role in the UK. Furthermore, export in regime 2 and regime 3 and narrow money in all regimes are the Granger causes of GDP based on the MS-Granger causality method. In addition the results of the traditional Granger causality test confirm the MS-Granger causality test results in all countries except from the UK's export.

Due to limited access to data for countries that may affect oil prices and the lack of seasonal variables in Iran, this paper focuses only on four countries. This may limit generalizability of the results. Accordingly, performing more research at the regional and international levels with a larger number of countries are recommended.

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Bank lending in an emerging economy: How does central bank reserve accumulation matter?

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Abstract

Purpose — The paper examines the impacts of the central bank's foreign exchange reserves on bank lending, captured by the dimensions of quantity (loan growth) and quality (credit risk).

Methods — This research analysis is based on bank-year observations in Vietnam during 2007–2019 and employs the two-step system Generalized Method of Moments in dynamic panels.

Findings — This study finds that banks tend to increase their loan growth rate in response to reserves accumulation. Banks also expand loans and cash items on their asset structure while subsequently slashing total security investments and disaggregate government bond holdings. Our results also indicate that the central bank reserves accumulation is associated with less credit risk and more financial stability of the banking system.

Implication — This paper supports the notion that reserve accumulation could be a complementary monetary policy tool for lending navigation and economic growth. Besides, reserve interventions may be used for financial stability, given the finding that it is found to curtail bank credit risk and financial instability.

Originality — This paper contributes to the literature by focusing on critical aspects of bank lending, including quantity and quality, to paint a bigger picture of the benefits and costs of reserve accumulation and decompose bank asset portfolios into disaggregate components, thereby providing more insight into bank responses.

Keywords — Bank lending, bank risk, central bank, foreign exchange reserves.

Introduction

Foreign exchange interventions relate to purchases (or sales) of foreign currency denominated assets and sales (purchases) of local currency assets, performed by central banks. In this regard, reserve accumulation could be seen as foreign savings financed by domestic borrowings of central banks. To neutralize the adverse impacts of foreign reserve purchases on money supply and interest rates in the market, central banks have to prevent excessive liquidity. It is established in the literature that most foreign exchange reserves were sterilized, with the sterilization instruments (commonly via the issuance of central bank securities) mainly held by domestic banks (Aizenman & Glick, 2009).

Foreign reserve accumulation may drive the safety and soundness of the financial system in multiple routes. When it comes to the bright side, with the precautionary motive, central banks

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hoard foreign reserves to guard against the likelihood of stuck international capital flows in future stressed times (Jeanne, 2007). It is found that during the global financial crisis, countries holding considerable foreign reserves may well circumvent exchange rate depreciations (Aizenman & Hutchison, 2012). Similarly, the central banks' foreign reserve holdings may improve market liquidity and cushion liquidity shocks that can hurt the financial market (Akdoğan, 2012). In a long-standing view, increased foreign exchange reserves allow the central bank to have wider and stronger interventions, thereby leaving the money supply unaffected and causing little effect on the exchange rate (Sarno & Taylor, 2001). Besides, one could usually argue that central banks engage in holdings of reserves for the mercantilist motive (Aizenman & Lee, 2007).

Another common literature stream emphasizes the dark side of foreign reserves. In this vein, the direct costs describe reserve holdings' financing costs (Adler & Mano, 2020). Also, indirect costs of reserve accumulation seem to be excessive. The over-accumulation of foreign reserves may initiate distortions in the market and wipe out the accumulation of other assets necessary for economic growth (Cook & Yetman, 2012). If reserves are extensively sterilized by increased reserve requirements or issued sterilization instruments purchased by banks, offsetting decisions made by these banks might hamper the financial system's overall riskiness. For instance, sterilization may result in persistent interest rate differentials and then persistent capital inflows, which is considered a contributing factor to the financial crisis (Cavoli & Rajan, 2006).

Despite the popularity of the reserve accumulation policy, the empirical literature has mainly focused on the macro benefit-cost tradeoff of reserve holdings, while the linkage between reserve accumulation and bank activities has been given little attention. This is surprising when commercial banks play a pivotal role in central banks' policy transmission, and the literature also implies that central banks' borrowings may induce profound implications in altering credit allocation and bank behaviors. For instance, banks might prefer to enhance their asset portfolios' risk levels to compensate for the influences of low-yield sterilization assets they are required to hold (Adler & Mano, 2020). In contrast, banks in markets with heavy sterilized reserves tend to hoard more safe sterilization securities than risky assets (Cook & Yetman, 2012). Hence, reserve accumulation may mitigate bank exposure to risk. Additionally, when central banks finance foreign reserve purchases by domestic borrowings, commercial banks are left with fewer funds and tend to cut loans granted to the real sectors (Yun, 2020). Being aware of the critical potential effects of reserve accumulation on the domestic banking system, we aim to fill this literature gap in the present study.

Despite the scarcity, some documents highlight the impacts of foreign reserves on bank investments, and our paper is relatively related to them. Through a panel of 20 Asian countries during 1980–2008, Lee and Choi (2010) display that foreign reserve accumulation, captured by the reserves-to-GDP ratio, may depress domestic investment rates. Cook and Yetman (2012) utilize a sample dataset of 55 Asian banks and conclude that bank loan growth declines in response to increased foreign exchange reserves. This result is repeated in Yun (2020)'s work, which demonstrates a crowding-out effect of reserve accumulation in Korea. Focusing on banks in Colombia, Hofmann, Shin, and Villamizar-Villegas (2019) reveal that foreign exchange reserves' sterilized buildup considerably alleviates new corporate lending.

This paper investigates the impacts of reserve accumulation on bank lending. Comprehensively, bank lending is captured by two dimensions of quantity (loan growth) and quality (credit risk). We also pay attention to components of bank asset portfolios, including bank loans, cash components, security investments, and even decomposed government bond holdings, to explore the sterilization practice. To deepen our findings, we analyze the heterogeneity effects across different banks according to their standard bank-specific characteristics. Specifically, the aspects of bank size, capitalization, and bank return are explored as inspired by the former literature, expecting to indicate whether different bank groups could be affected differently by central bank reserve accumulation. We employ the two-step system generalized method of moments (GMM) in dynamic panels with bank-level data during 2007–2019 for Vietnam, an emerging market that has grown rapidly over the years, to display evidence for the impacts of reserve accumulation. Vietnam offers a favorable case to test our research issues. Vietnam's economic growth is largely financed by bank lending, which is considered a core economic indicator (Dang & Dang, 2020). Vietnam appears to be open to capital inflows and targeting exchange rates through foreign exchange interventions has been regularly used here. The scale of foreign exchange reserves has changed and grown dramatically in Vietnam over the past decade (see section 3). The State Bank of Vietnam (SBV) usually declared that it had sterilized most foreign exchange purchases to absorb excessive liquidity. Besides, the Vietnamese banking system is also featured by various reforms recently, which leads to significant changes in the industry with high differentiation among participants (Dang & Huynh, 2020).

Sharing a similar topic with the works mentioned above, however, our paper with a different setting and startling results still adds important contributions to the literature. We focus on two key aspects of bank lending, including quantity and quality, to paint a bigger picture of the benefits and costs of reserve accumulation, which previous studies have not exploited. We also decompose bank asset portfolios into disaggregate components to offer more insight into bank responses. In this regard, our essential contribution to the current debate is to shed light on the expansionary effects of reserve accumulation on credit conditions so that foreign exchange intervention could complement monetary policy in credit supply, economic growth, and financial stability. The main findings of this study challenge those obtained previously in other markets and thus balance the understanding of the topic. Notably, the present paper offers in-depth conclusions by exploiting the heterogeneity effects across different banks; this is a special note as well as an innovation of this study.

Methods

Estimation Strategies

To explore the impact of reserve accumulation on bank lending in Vietnam, we specify the model equation as follows:

$$Y_{i,t} = a_0 + a_1 \times Y_{i,t-1} + a_2 \times FXR_{t-1} + a_3 \times X_{i,t-1} + a_4 \times Z_{t-1} + \varepsilon_{i,t} \quad (1)$$

where i and t express banks and years, respectively. The dependent variable Y is either the quantity of bank lending, captured by the growth rate of loans, or the quality of bank lending, measured by the loan loss provision ratio and the non-performing loan ratio. The lagged dependent variable is included in the right-hand side of the specification to adopt the dynamic feature in bank lending. FXR is the measure of reserve accumulation, while X contains bank-specific controls, and Z is a vector of macroeconomic variables. $\varepsilon_{i,t}$ is the error term. The lags of independent variables are preferred as bank lending may not react immediately to changes in both macro- and micro-events.

As a straight-forward approach, we use the natural logarithm of the SBV's foreign exchange reserves (*lnFXR*) to reflect our primary explanatory variable (reserve accumulation exhibited in US dollars). In addition to this baseline variable, we also employ an additional proxy allowing for the size of the economy and the floating/fixed exchange rate regime (Chen, Wu, Jeon, & Wang, 2017). The equation to compute the alternative proxy for reserve accumulation is written as follows:

$$FXR/GDP = \left(\frac{fxr}{GDP}\right) \times \frac{\sigma(fxr)}{\sigma(fxr) + \sigma(eer)} \quad (2)$$

where *fxr* represents the SBV's foreign exchange reserves, *GDP* captures the gross domestic products, $\sigma(fxr)$ and $\sigma(eer)$ indicate the standard deviations of monthly indexed foreign exchange reserves and effective exchange rates, respectively. The identical estimation results from our two explanatory variables of main interest as set up above would signify the robustness of our findings.

To draw a comprehensive picture of reserve accumulation's impact on bank activities, we utilize an additional measure to display bank financial stability. This measure has been widely employed in the banking literature in the form of the Z-score index (Ngambou Djatche, 2019; Niu, 2012), via the formula specified as:

$$Z = \frac{ROA + capital}{\sigma(ROA)} \quad (3)$$

where *ROA* denotes return on assets, *capital* represents the capital ratio and *σ(ROA)* indicates the standard deviation of *ROA* over the entire period under research. Inspired by the former literature, we take the natural logarithm of the Z-score index in the regression stage (Ngambou Djatche, 2019).

Additionally, to control for other potential factors that may drive loan growth and bank risk, we incorporate some standard variables commonly suggested by the existing literature (Dang & Dang, 2020; Delis & Kouretas, 2011; Khan, Ahmed, & Gee, 2016; Yang & Shao, 2016). For bank-specific control variables, we include bank size, capital equity, and bank return; for macroeconomic factors, we include inflation and economic cycle. Specific definitions of all variables are given in Table 1.

With the presence of the lagged dependent variable among our independent variables, we employ the GMM estimator to perform regressions. This GMM estimator is also suited for dealing with omitted variables, reverse causality, and measurement errors (Arellano & Bover, 1995; Blundell & Bond, 1998). Following Roodman (2009), we apply the procedure to curb the proliferation of instruments used for small samples. To assure the consistency of the GMM estimation, we perform the Hansen test of valid instruments and the AR(1)/AR(2) tests for the first- and second-order serial correlation.

Data

We collect bank-level data from the financial reports published by Vietnamese commercial banks for the period 2007–2019. According to the banking regulation of Vietnam, the structure of commercial banks' financial reports has been strictly standardized in a consistent form since 2007, which has to be audited. Moreover, before 2007, only a few large banks fully published their financial reports. We eliminate banks that provide insufficient information for the calculation of required variables. For the macroeconomic data, we source the foreign exchange reserves from the International Financial Statistics (IFS), while the GDP growth rate and the inflation rate are obtained from the World Development Indicators (WDI). Our final sample covers 31 Vietnamese banks and constitutes an unbalanced panel with 391 bank-year observations, making up on average 90% of the Vietnamese banking system's total assets.

Results and Discussion

Preliminary Results

Table 1 reports the descriptive statistics for all variables. In general, through the values of standard deviations and the spreads of extreme values, we can observe a wide dispersion across banks in the working and outcome of lending activities.

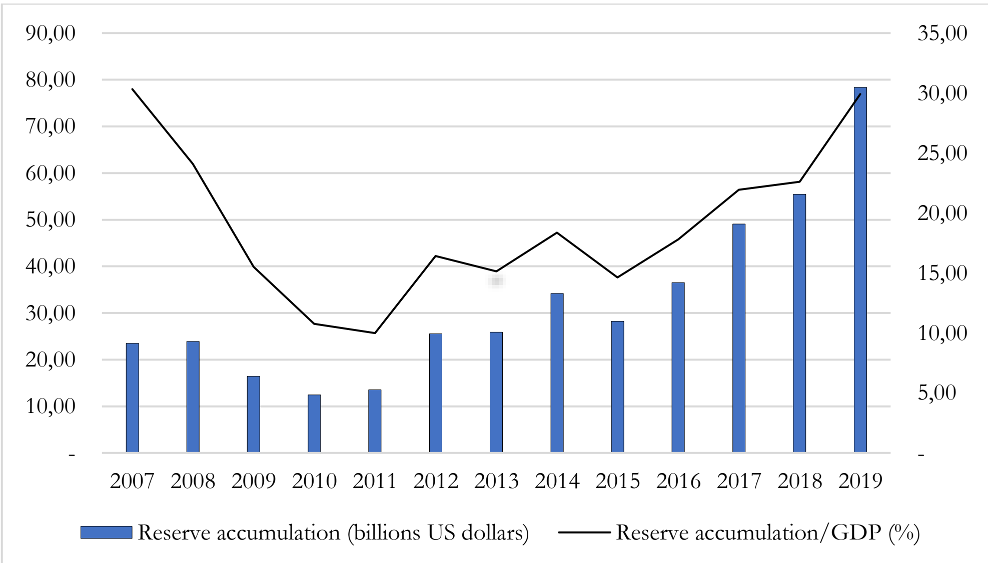

Figure 1. The evolution of the average reserve accumulation over the sample period of 2007–2019

Regarding the evolution of the SBV's foreign exchange reserves, Figure 1 depicts the average reserve accumulation over our sample period from 2007 to 2019. After a period of alternate downward and upward trends from 2007 to 2014, since 2015, we could observe substantial increases in overall central bank reserves. The increases led to the fact that Vietnam's reserve accumulation almost tripled from 23.5 billions USD (2007) to 78.3 billions USD (2019). Compared to the SBV's foreign exchange reserve scale, the reserve accumulation relative to Vietnam's GDP displayed a slightly different pattern. Experiencing a period of significant fluctuations up and down, which highlights the U-shape pattern, the rate in 2019 (29.91%) was almost the same as in 2007 (30.33%). Though, we still observe the continuous increase of this rate since 2015, similar to the reserve scale. Bearing in mind the difference in the changes of the relative and absolute foreign exchange reserves, we expect the identical results in the regression section from the two main explanatory variables could highlight the robustness of our findings.

Table 1. Definitions and summary statistics of variables employed

	Mean	SD	Min	Max	Definitions
Bank lending and investment variables					
Loangrowth	29.533	29.671	−5.159	111.120	The annual growth rate of bank loans (%)
Loanshare	54.724	12.645	31.227	74.392	The ratio of bank loans to total assets (%)
Cash	17.442	9.603	5.570	38.193	The ratio of liquid assets (cash plus due from other institutions) to total assets (%)
Securities	16.813	7.364	4.470	31.980	The ratio of security investments to total assets (%)
Governbonds	8.615	4.696	1.606	18.851	The ratio of government bond holdings to total assets (%)
Bank risk measures					
lnZscore	2.867	0.488	1.915	3.662	Natural logarithm of the Z-score index
LLP	1.253	0.509	0.502	2.499	The ratio of loan loss provisions to total gross loans (%)
NPL	2.147	1.187	0.495	5.159	The ratio of non-performing loans to total gross loans (%)
Bank-level characteristics					
Capital	10.072	4.647	4.939	21.884	The ratio of bank equity to total assets (%)
Size	31.972	1.233	29.943	34.269	Natural logarithm of total assets
ROA	0.911	0.656	0.038	2.279	The ratio of net return on total assets (%)
Macroeconomic factors					
lnFXR	24.074	0.521	23.246	25.084	Natural logarithm of the SBV' foreign exchange reserves
FXR/GDP	17.498	6.361	7.632	29.460	The SBV' foreign exchange reserves as a fraction of GDP, taking into account the exchange rate regime (%)
GDP	6.245	0.642	5.247	7.130	The annual growth rate of GDP (%)
Inflation	7.495	6.226	0.631	23.115	The annual inflation rate (%)

The sample covers 31 commercial banks in Vietnam over the period 2007–2019, with a total number of 391 observations, forming an unbalanced panel data.

We now move on to the estimation parts obtained using the consistent GMM estimator, which is validated by the Hansen test results of valid instruments and the AR(1)/AR(2) tests for the first second-order serial correlation. Some reliable and interesting findings have emerged.

Central Bank Reserve Accumulation and Loan Growth

We report the baseline estimation results for the linkage between reserve accumulation and bank loan growth in Table 2. Both columns 1 and 2 presenting the estimation results for alternative reserve accumulation variables indicate that the coefficient of reserve accumulation is significantly positive at the 1% level. This result shows that reserve accumulation induces a significant positive effect on bank lending growth in Vietnam; in other words, we find that the overall bank loan growth rate increased significantly after reserve accumulation from the central bank.

Quantitatively, our results are also economically significant. Based on the coefficient of column 1 (Table 2), we could infer that an increase of one standard deviation in the reserve accumulation (0.521) is associated with an increase in the loan growth rate of 6.53 percentage points. Such an effect is reasonable given the mean value of the loan growth rate is 29.53%.

Table 2. Reserve accumulation and bank loan growth

	(1) Loangrowth	(2) Loangrowth
Lag of Loangrowth	0.199***	0.109***
	(0.014)	(0.015)
lnFXR	12.539***	
	(0.924)	
FXR/GDP		0.623***
		(0.049)
Capital	1.107***	1.088***
	(0.173)	(0.168)
Size	−1.804***	−1.605**
	(0.610)	(0.665)
ROA	−2.747***	−2.999***
	(0.938)	(1.116)
GDP	−7.473***	−7.161***
	(0.508)	(0.609)
Inflation	1.023***	0.705***
	(0.082)	(0.089)
Observations	360	360
Banks	31	31
Instruments	29	29
AR(1) test	0.001	0.001
AR(2) test	0.168	0.120
Hansen test	0.149	0.144

Notes: The dependent variable is the growth rate of bank loans. Symbols *** and ** indicate significance at the 1% and 5% levels, respectively.

Table 3. Reserve accumulation and bank portfolio components (using the natural logarithm of foreign exchange reserves)

	(1) Loanshare	(2) Cash	(3) Securities	(4) Governbonds
Lag of Loanshare	0.604***			
	(0.030)			
Lag of Cash		0.420***		
		(0.042)		
Lag of Securities			0.692***	
			(0.035)	
Lag of Governbonds				0.656***
				(0.043)
lnFXR	3.204***	2.362***	−1.557***	−1.103***
	(0.642)	(0.512)	(0.223)	(0.172)
Capital	−0.030	−0.126*	0.025	−0.069***
	(0.050)	(0.069)	(0.040)	(0.024)
Size	2.008***	−1.792***	−0.353***	0.090
	(0.342)	(0.339)	(0.134)	(0.057)
ROA	−0.238	7.465***	−0.675***	−0.972***
	(0.310)	(0.926)	(0.138)	(0.151)
GDP	1.031***	−0.414*	−1.882***	−0.450***
	(0.188)	(0.234)	(0.262)	(0.116)
Inflation	0.199***	−0.112**	−0.166***	−0.049***
	(0.030)	(0.051)	(0.019)	(0.010)
Observations	360	360	360	360
Banks	31	31	31	31
Instruments	29	29	29	29
AR(1) test	0.001	0.000	0.000	0.001
AR(2) test	0.436	0.915	0.443	0.779
Hansen test	0.281	0.286	0.161	0.336

Notes: The dependent variables are loans, cash and due, total security investment, and sovereign bond holdings, divided by total assets (the variables of Loanshare, Cash, Securities, and Governbonds are exhibited respectively at the top of each column). Symbols ***, ** and * indicate significance at the 1%, 5%, and 10% levels, respectively.

To deepen our findings, we pay attention to the components of bank asset portfolios. We now turn to use new indicators as the dependent variables, including bank loans, cash components (cash and due from other institutions), and security investments, all computed as a share of total assets. We also decompose the security investments to take government bond holdings to explore the sterilization conducted and absorbed. We redo our estimations with the new dependent variables using the system GMM estimator and report all results in Tables 3 – 4.

Table 4. Reserve accumulation and bank portfolio components (using the foreign exchange reserves as a fraction of GDP)

	(1) Loanshare	(2) Cash	(3) Securities	(4) Governbonds
Lag of Loanshare	0.559***			
	(0.024)			
Lag of Cash		0.549***		
		(0.027)		
Lag of Securities			0.678***	
			(0.037)	
Lag of Governbonds				0.651***
				(0.045)
FXR/GDP	0.148***	0.103***	−0.059***	−0.053***
	(0.031)	(0.015)	(0.021)	(0.013)
Capital	0.058	0.063	0.003	−0.084***
	(0.052)	(0.052)	(0.043)	(0.025)
Size	2.238***	−1.031***	−0.548***	−0.036
	(0.294)	(0.162)	(0.113)	(0.063)
ROA	−1.184***	1.960***	−0.378**	−0.731***
	(0.265)	(0.221)	(0.150)	(0.122)
GDP	1.398***	0.178	−1.992***	−0.509***
	(0.189)	(0.159)	(0.281)	(0.133)
Inflation	0.128***	−0.074**	−0.149***	−0.036***
	(0.032)	(0.035)	(0.016)	(0.010)
Observations	360	360	360	360
Banks	31	31	31	31
Instruments	29	29	29	29
AR(1) test	0.001	0.000	0.000	0.002
AR(2) test	0.624	0.973	0.498	0.751
Hansen test	0.241	0.220	0.145	0.320

Notes: The dependent variables are loans, cash and due, total security investments, and sovereign bond holdings, divided by total assets (the variables of Loanshare, Cash, Securities, and Governbonds are exhibited respectively at the top of each column). Symbols *** and ** indicate significance at the 1% and 5% levels, respectively.

We notice that the coefficients on accumulated reserves are significantly positive in the function of bank loans and cash share but significantly negative in the equation of total securities and disaggregate sovereign bonds holdings. This set of results implies that in response to the accumulation of foreign exchange from the central bank, the proportion of loans and cash on the bank asset structure tends to be enlarged, in contrast to security investments. Interestingly, a deeper pattern reveals that banks have decreased the holdings of government bonds subsequently.

Our findings strongly challenge those obtained previously in other markets, highlighting another cost of foreign exchange reserves in decreasing bank credits to the real sectors (Cook & Yetman, 2012; Hofmann et al., 2019; Yun, 2020). In turn, we lend support to the work of Chen et al. (2017), which posits that unless thoroughly sterilized, foreign exchange intervention may modify the monetary base and play a role similar to monetary policy. Overall, some mechanisms could be used to explain our main findings so far. Accumulation of foreign reserves by the central bank may enhance bank liquidity positions. According to the "portfolio balance" channel (Tobin, 1969), banks do not consider cash as a perfect substitute for the assets they sell to the central bank. Thus, they are incentivized to use the cash received to invest in high-return assets. To avoid excessive liquidity in the economy, the central bank is supposed to sterilize the reserves accumulation. However, sterilization may be incomplete. For instance, if sterilization bills and money share the feature of close substitutes, then more holdings of sterilization bills (mostly in the form of government bonds via open market operations in Vietnam) are not ideal to accepted by banks, and it may not offset the expansionary influences of foreign exchange reserves (Cook & Yetman, 2012). Hence, banks may choose to invest more in higher-yield items such as loans or possibly keep their money received in cash and deposits at other institutions.

Central Bank Reserve Accumulation and Bank Risk

Table 5 reports the estimation results for how central bank reserve accumulation drives bank riskiness, captured by the measures of credit risk and overall risk (or bank financial stability). We document that the coefficients associated with loan loss provisions and non-performing loans are significantly negative at the 1% level, while those associated with the Z-score measure are significantly positive at the 1% level. These results demonstrate that banks suffer less credit risk and become more financially stable after the central bank accumulates more foreign exchange reserves.

In some detail, using the face value of coefficients in column 2 (Table 5), we could calculate that an increase of one standard deviation in reserve accumulation may cause a decline of 0.42 percentage points in the non-performing loan ratio. Simultaneously, the coefficients in column 6 reveal that an increase of one percentage point in the reserve accumulation relative to GDP is associated with a rise of 0.002 percentage points in the measure of bank financial stability. The magnitudes of coefficients all outline the economic plausibility of our findings.

Table 5. Reserve accumulation and bank riskiness

	(1) LLP	(2) NPL	(3) lnZscore	(4) LLP	(5) NPL	(6) lnZscore
Lag of LLP	0.579***			0.664***		
	(0.032)			(0.032)		
Lag of NPL		0.540***			0.547***	
		(0.035)			(0.026)	
Lag of lnZscore			0.979***			1.005***
			(0.018)			(0.027)
lnFXR	−0.253***	−0.801***	0.032***			
	(0.016)	(0.065)	(0.006)			
FXR/GDP				−0.014***	−0.037***	0.002***
				(0.002)	(0.003)	(0.001)
Capital	−0.015***	−0.017		−0.020***	−0.024***	
	(0.003)	(0.011)		(0.002)	(0.009)	
Size	0.027*	−0.110***	0.036***	−0.039***	−0.188***	0.038***
	(0.014)	(0.035)	(0.005)	(0.011)	(0.028)	(0.005)
ROA	−0.002	−0.223***		0.050***	−0.094**	
	(0.017)	(0.047)		(0.017)	(0.038)	
GDP	0.055***	0.290***	0.059***	0.073***	0.276***	0.054***
	(0.008)	(0.034)	(0.007)	(0.014)	(0.029)	(0.010)
Inflation	0.006***	−0.002	−0.001	0.007***	0.006	−0.002*
	(0.002)	(0.006)	(0.001)	(0.002)	(0.005)	(0.001)
Observations	360	360	360	360	360	360
Banks	31	31	31	31	31	31
Instruments	29	29	27	29	29	27
AR(1) test	0.000	0.000	0.001	0.000	0.000	0.001
AR(2) test	0.241	0.101	0.509	0.257	0.106	0.518
Hansen test	0.291	0.155	0.162	0.387	0.153	0.150

Notes: The dependent variables are the loan loss provision ratio, the non-performing loan ratio, and the natural logarithm of the Z-score index (the variables of LLP, NPL, and lnZscore are exhibited respectively at the top of each column). Symbols ***, ** and * indicate significance at the 1%, 5%, and 10% levels, respectively. When performing estimations in the Z-score measure model, we exclude capital and return variables to avoid spurious regressions since these variables are the main factors in computing the Z-score index.

Generally, we gain strong evidence to prove that reserve accumulation has protected the Vietnamese banking system from credit risk and financial instability. Although banks do not show absolute improvement in the underlying liquidity position after large-scale reserve accumulation (holdings of money increase but securities decrease), non-easy gains from holdings of government securities (as sterilization instruments) may encourage banks to become more efficient in other operations (Mohanty & Turner, 2006), especially credit segments. This argument is entirely consistent with our finding obtained previously. Furthermore, increasing foreign exchange reserves could enable the central bank to have broader and more powerful interventions, thereby leaving the money supply unaffected and inducing little effect on the exchange rate (Sarno & Taylor, 2001). This long-standing view indicates safer asset portfolios of banks.

Heterogeneity Effects

In this subsection, we try to offer more insight into our findings by identifying the effects of reserve accumulation on bank lending while taking into account differences across banks. The literature inspires our strategy on monetary policy transmission via the bank lending channel and the bank risk-taking channel. Accordingly, these literature strands have shown that monetary shocks exert stronger impacts on banks that are more vulnerable (i.e., smaller banks, lower-capitalized banks, or less liquid banks that have limited access to non-deposit funding) (Delis & Kouretas, 2011; Khan et al., 2016; Yang & Shao, 2016). If similar mechanisms still work when financial conditions are adjusted generally, central bank reserve accumulation that relaxes money supply would be expected to produce more pronounced effects on financially weaker banks.

Table 6. Heterogeneity in the impact of reserve accumulation on bank loan growth

	The dependent variable is the growth rate of bank loans					
	(1)	(2)	(3)	(4)	(5)	(6)
Lag of Loangrowth	0.152***	0.169***	0.187***	0.154***	0.094***	0.151***
	(0.019)	(0.015)	(0.018)	(0.020)	(0.020)	(0.022)
lnFXR	11.083***	16.105***	10.113***			
	(0.964)	(1.908)	(1.377)			
lnFXR*Capital	−0.089***					
	(0.007)					
lnFXR*Size		−0.155**				
		(0.064)				
lnFXR*ROA			−0.651***			
			(0.096)			
FXR/GDP				2.290***	0.700***	1.701***
				(0.178)	(0.059)	(0.184)
FXR/GDP*Capital				−0.220***		
				(0.021)		
FXR/GDP*Size					−0.005**	
					(0.002)	
FXR/GDP*ROA						−0.829***
						(0.133)
Capital	2.118***	1.091***	0.670***	2.497***	1.029***	1.131***
	(0.157)	(0.173)	(0.163)	(0.170)	(0.166)	(0.247)
Size	−4.063***	1.654	−2.968***	−5.590***	−1.574***	−1.292
	(0.484)	(2.206)	(0.629)	(1.081)	(0.546)	(0.859)
ROA	−0.745	−2.325**	−9.514***	−0.603	−3.319**	9.367**
	(1.085)	(1.059)	(1.191)	(1.433)	(1.292)	(3.654)
GDP	−5.997***	−6.577*	−9.217***	−4.987***	−6.284***	−7.117***
	(0.509)	(0.529)	(1.116)	(0.664)	(0.546)	(0.634)
Inflation	0.982***	0.929***	0.738***	0.862***	0.696***	0.787***
	(0.089)	(0.112)	(0.130)	(0.088)	(0.088)	(0.062)
Observations	360	360	360	360	360	360
Banks	31	31	31	31	31	31
Instruments	30	30	30	30	30	30
AR(1) test	0.001	0.001	0.000	0.001	0.001	0.001
AR(2) test	0.272	0.171	0.198	0.525	0.119	0.142
Hansen test	0.154	0.160	0.146	0.174	0.151	0.131

Notes: Symbols *** and ** indicate significance at the 1% and 5% levels, respectively.

Aligning our present paper to the former literature, we explore the following bank-specific characteristics that can modify the linkage between reserve accumulation and bank lending: bank capital, bank size, and bank return. Regarding the econometric approach, we extend the baseline model by interacting reserve accumulation variables with different bank-specific characteristics separately. The interaction terms allow us to verify the heterogeneity in how bank lending reacts to reserve accumulation. The estimation results for the function of bank loan growth and bank risk (only captured by the loan loss provision ratio for the sake of brevity, while other variables still yield the same results) are presented in Tables 6–7.

Table 7. Heterogeneity in the impact of reserve accumulation on bank risk

	The dependent variable is the ratio of loan loss provisions to total gross loans					
	(1)	(2)	(3)	(4)	(5)	(6)
Lag of LLP	0.561***	0.569***	0.623***	0.673***	0.670***	0.665***
	(0.035)	(0.031)	(0.032)	(0.040)	(0.029)	(0.034)
lnFXR	−0.262***	−0.356***	−0.251***			
	(0.016)	(0.037)	(0.014)			
lnFXR*Capital	0.001***					
	(0.000)					
lnFXR*Size		0.003***				
		(0.001)				
lnFXR*ROA			0.006***			
			(0.001)			
FXR/GDP				−0.012***	−0.150***	−0.014***
				(0.002)	(0.025)	(0.002)
FXR/GDP*Capital				0.001***		
				(0.000)		
FXR/GDP*Size					0.004***	
					(0.001)	
FXR/GDP*ROA						0.001
						(0.001)
Capital	−0.019***	−0.017***	−0.020***	−0.016***	−0.021***	−0.020***
	(0.003)	(0.003)	(0.003)	(0.003)	(0.002)	(0.003)
Size	0.047***	−0.065**	0.009	−0.047***	−0.113***	−0.039***
	(0.015)	(0.030)	(0.012)	(0.010)	(0.023)	(0.014)
ROA	0.004	−0.001	−0.055***	0.056***	0.050***	0.045***
	(0.016)	(0.016)	(0.019)	(0.017)	(0.018)	(0.015)
GDP	0.054***	0.052***	0.052***	0.104***	0.069***	0.071***
	(0.009)	(0.009)	(0.010)	(0.013)	(0.012)	(0.016)
Inflation	0.007***	0.007***	0.006***	0.008***	0.008***	0.007***
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
Observations	360	360	360	360	360	360
Banks	31	31	31	31	31	31
Instruments	30	30	30	30	30	30
AR(1) test	0.000	0.000	0.000	0.000	0.000	0.000
AR(2) test	0.217	0.254	0.405	0.265	0.292	0.262
Hansen test	0.356	0.449	0.364	0.315	0.344	0.378

Notes: Symbols *** and ** indicate significance at the 1% and 5% levels, respectively.

For both measures of reserve accumulation, we find that the regression coefficients on interaction terms across different aspects of bank balance sheets are significantly negative in the equation of loan growth (Table 6) and significantly positive in the bank risk specification (Table 7). All signs of these coefficients are opposite to those of stand-alone reserve accumulation variables. These results suggest that the impacts of central bank reserve accumulation on bank lending, both in dimensions of quantity and quality, vary according to bank-specific characteristics. More precisely, weaker banks (i.e., smaller banks, lower-capitalized banks, and less profitable banks) are more sensitive to the impacts of reserve accumulation on bank lending.

On the one hand, in case of reserve accumulation, financially stronger banks have incentives as well as are encouraged to cooperate with the central bank on sterilization. If these stronger banks take over more sterilization securities, they will be less able to increase loans. In this scenario, more sterilized intervention could also modify bank behavior in the mechanism that easy profits from holdings of sterilization securities could weaken banks' pressures to become more efficient in monitoring and supervising loans (Mohanty & Turner, 2006). On the other hand, foreign reserve decumulation of the central bank leads to a fall in the volume of loanable funds and may subsequently contract lending of banks. Such a contraction may be more pronounced at weaker banks, as they find it more challenging and expensive to reach external funds to finance their lending. Moreover, reserve decumulation is typically done in periods of extended uncertainty (Yun, 2020), when weaker banks are more vulnerable. This situation makes the riskiness of these banks increase to a higher level.

Conclusion

The study examines the impacts of central bank reserve accumulation on lending activities in the banking industry, captured by two dimensions of quantity (loan growth) and quality (credit risk). Using bank-level data and foreign exchange reserves during 2007–2019 in Vietnam — an emerging market that has tripled its reserves over the past decade, we document multiple interesting findings. In response to the reserve accumulation, banks increase their loan growth rate. Our additional analysis reveals that loans and cash items on the bank asset structure tend to be expanded in this vein, while total security investments and the holdings of government bonds may be narrowed subsequently. Our results also indicate that foreign reserves drive bank risk profiles in the way that banks suffer less credit risk after the reserve accumulation of the central bank. Focusing on a broader aspect of bank riskiness, we further find that bank financial stability tends to be improved when the central bank in Vietnam increases its foreign exchange reserves. Most interestingly, as a special note to the findings of this study, our results reveal that weaker banks (i.e., smaller banks, lower-capitalized banks, and less profitable banks) are more responsive to how central bank reserve accumulation affects bank lending. This pattern firmly holds for the dimensions of loan growth and credit risk across alternative measures of reserve accumulation.

This paper offers some policy implications. The empirical evidence presented displays that reserve accumulation has expansionary effects, i.e., banks increase loans to the economy and hold less risk-free securities. This supports the notion that reserve accumulation could be a complementary monetary policy tool for lending navigation and economic growth. Besides, reserve interventions may be used for the purpose of financial stability, given the finding that it is found to curtail bank credit risk and financial instability. However, monetary authorities should be concerned about the dark side of excessive liquidity after accumulating foreign exchange reserves since our findings cast some doubt on the effectiveness of the sterilization operations in Vietnam. Additionally, our study also calls for policy measures that pay attention to heterogeneous responses of different bank groups to central bank reserve accumulation.

We acknowledge that our paper is refined to a small single country with a limited sample size using yearly observations. We expect our work could be expanded to other countries/regions, which may enrich the current knowledge. Besides, we realize that there are several potential channels through which central bank reserves can alter bank lending's working and efficiency, albeit it is challenging to be sure about their significance based on empirical evidence in the present research. Hence, we leave the more rigorous empirical analysis of precise channels to future research to fully explain the impact of central bank reserves on bank lending.

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Similarity evidence between the country risk and the idiosyncratic risk: An empirical study of the Brazilian case

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Abstract

Purpose — This paper estimates the idiosyncratic risk (IDR) time series in the Brazilian economy and verifies its interaction with the Brazilian country risk indicators, measured by the EMBI+ (the Emerging Markets Bond Index).

Methods — This paper estimates various regression models to capture the dynamic nature of the variables. The models include the heteroscedastic conditional autoregressive models and vector error correction models (VECM).

Findings — The results show similarities or associations between the two indicators with interactions in the short and long run. The idiosyncratic risk proves to be a relevant indicator of the risk of economic activities implemented within the scope of the Brazilian economy and can help evaluate investments in related projects. This results also provide evidence of cointegration between the EMBI+ and IDR variations.

Implication — This result suggests an alternative way for obtaining estimates of the expected return required by economic agents in financing and investing in productive and infrastructure projects necessary for developing the Brazilian economy that provides greater employability and good social welfare.

Originality — This paper provides an alternative estimate of the time series proxy of idiosyncratic risk in the Brazilian economy. It also compares the results with the time series results obtained from the country risk measure EMBI+, widely used among resource managers in the international markets.

Keywords — Idiosyncratic risk, country risk, causality, cointegration

Introduction

Knowledge of available indicators is necessary for decision-making in investing and financing the productive projects and the trade-off risk-return for an optimal allocation of resources in an economy. It is essential in, especially, financial resources shortage periods, one frequent situation in many developing economies and, particularly, in the Brazilian economy.

Economic policymakers and managers in emerging economies attempt to attract capital to productive projects that develop their economies by providing greater employment opportunities and social welfare. On the other hand, resource managers seek to invest and finance projects or economic activities that can offer returns that compensate the risk assumed by allocating resources in these economies as well as any investor attempt to obtain the most meaningful information about their investment possibilities among investment options available in the international market.

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Investment and financing of productive projects or the production in economies occur through the financial market, particularly the capital market. The productive projects provide the necessary conditions for the economic development of national economies and to maintain jobs and income growth in the capital market, the source of resources for investment and financing. Therefore, a relevant advanced indicator of an economy is its capital market performance, whose prime indicator is the stock market profitability index that reflects the expectations of economic agents. From these stock market profitability indices, risks associated with productive projects, national economies sectors, a national economy, and the global economy can be estimated. The total risk of asset projects portfolio and its two components, the market risk and the idiosyncratic or specific risk can be estimated through the assets or projects portfolio's returns and the market portfolio returns. The finance theory studies investment and financing through appropriate methodologies to observe the risk associated with productive activities or productive projects. This theory was developed in the last seven decades from the portfolio optimization model, suggested by Markowitz (1952), and the model's determination specifies the risk-return trade-off. Hence, the risk quantification of a diversified portfolio can occur through the market model, presented in the literature by Sharpe (1963) in one of the finance theory seminal works. The market model can be represented by a stochastic model with an asset or an asset portfolio return as a response variable and the market portfolio return as a regressor whose proxy can be obtained in each national economy's stock market profitability indexes.

This way, each national economy provides an indicator representing the market portfolio that encompasses the most productive projects: the indicator closest to a market portfolio of that economy. It can be represented by the stock market profitability index of each of these economies. In contrast, the indicator representing the global market portfolio or the world economy can be obtained by a portfolio that gathers all national markets or a profitability index of that portfolio or the global stock market. Some financial market institutions calculate and make profitability indices for the national stock markets and the global stock market available. This way, one can build market models for each national economy and, in particular, for the Brazilian economy.

The total risk of each economy integrated into the world economy can be estimated as well as their components, the market risk, and the idiosyncratic risk. The market risk reflects the events that affect each national economy, such as wars, climatic disasters, and pandemics. In contrast, the idiosyncratic risk is associated with events that affect only a particular national economy or a country, that is, the risk part that may produce a direct impact motivated by a specific economic measure that affects a national economy. Therefore, the expectations of idiosyncratic risk indicate how each economy is mainly affected by events that concern only a country or how each country manages its productive projects and economic activity inherent problems. Hence, idiosyncratic risk can be recognized as a relevant indicator for investors in general.

Another relevant parameter associated with the risk of economic activity in national economies concerns the sovereign risk premium measured by country risk, whose main indicator is the Emerging Markets Bond Index (EMBI+). This index quantifies the margin that international investors consider a fair return for investing in productive projects or investing in bonds referenced by productive projects developed in a given national economy.

Many types of research have been carried out on the idiosyncratic risk of investment projects, investment project portfolios, or national economies in general. These studies are differentiated by how this risk is estimated, such as the models used and estimation methods, and the relationship or association of the indicators. It happens with the returns or with other indicators of these investment projects or investment project portfolios. The same occurs in researches related to sovereign risk or the premium associated with this risk. Further on, in this work, some of these researches will be commented.

This work aims to estimate the idiosyncratic risk time series of the Brazilian economy and verify its interaction with the EMBI+ Brazil, the Brazilian country risk indicator. It does so using heteroscedastic conditional autoregressive models, autoregressive vector models with error correction, respectively. This paper intends to provide an indicator for the risk associated with the Brazilian economy as an alternative to country risk.

In addition to this introduction, other sections are presented in this work as follows. Section 2 presents a brief of the research related to the topics discussed here, while Section 3 presents its methodological approach. The sample or data employed that are presented in Section 4. And the analysis of the results obtained is presented in Section 5. Finally, in Section 6 are the conclusion and final comments of the work followed by the bibliographic references.

The stock profitability index is a proxy for a diversified portfolio for each of the economies that make up the global economy, which allows estimating national and global economy volatilities. Asset volatility or total risk of investment projects portfolios or an economy can be divided into two parts. The first one is the market risk, which refers to the risk affecting the portfolio or all economies inserted in the global market. The second one is the idiosyncratic risk, which can be minimized with the diversification of the portfolio and recognized as an inherent risk to each national economy inserted in the global economy. Idiosyncratic risk expectations describe how each economy is particularly affected by events related to only a respective country. Each country manages the problems related to its products and infrastructure projects that can provide well-being.

From Sharpe (1963), many types of research have been carried out to determine the total risk and its components. The market model estimation allows for determining the beta coefficient, the primary indicator of market risk, and the idiosyncratic risk also called specific risk. In a pioneering work, Rosenberg and McKibben (1973) search to make equity systematic and specific risk predictions proposing stochastic models construction using the market model concept to determine the beta coefficient and the idiosyncratic. Based on the work of Rosenberg and McKibben (1973)-many types of research were developed to extend and improve the methods employed. Fu (2009) sought to estimate the idiosyncratic risk and its relationship with the monthly stock returns using the three-factor model suggested by Fama and French (1993) and the Autoregressive Conditional Heteroscedasticity (ARCH) models family available in the finance literature, developed from the seminal work by Engle (1982) and the work by Bollerslev (1986). Angelidis and Tessaromatis (2008) observe that idiosyncratic risk has been neglected to the detriment of a more considerable emphasis given to the systematic risk in determining the risk premium, and Campbell, Hilscher, and Szilagyi (2008) emphasize the importance of assessing idiosyncratic risk.

In general, market volatility increases in the periods leading up to crises and during periods of crisis. In the recent period of the world economy and as noted by Kalra (2008) with the North American subprime crisis, which started in the middle of 2007, and the decrease in the global market credit supply: the global economy and in capital markets, in particular, was out of control. Among the most recent studies that attempt to estimate idiosyncratic risk are Chang, Ko, Nakano, and Ghon Rhee (2018), with data from the Japanese economy, Shi and Zhou (2019), with data from the Chinese stock market, and Blitz, Hanauer, and Vidojevic (2020), which treats this risk as an anomaly, can be mentioned. Works related to sanitary crises or epidemics and pandemics that have occupied academic researchers with studies that seek to verify changes in the productive projects, in productive projects portfolios, in economic sectors, and national economies risk, should also be mentioned. Among these researches, Naidenova, Parshakov, and Shakina (2020), which deals with idiosyncratic and systematic shocks caused by the Covid-19 Pandemic in the financial markets, and the recent research by Atkenson (2020), by Barro, Ursua, and Weng (2020), by Anderson, Heesterbeek, Klinkenberg, and Hollingsworth (2020), Mckibbin and Fernando (2020), and Gourinchas (2020), on the influence of the crisis caused by the Covid-19 Pandemic in the macroeconomic environment.

A prominent aspect of emerging economies is the level of accentuated indebtedness, in particular external indebtedness. Since in several periods, there is a scarcity of resources requiring external resources contribution. Hence, from the resource managers' point of view, all available indicators are crucial for efficient allocation of resources in these economies and assisting investment decision making. Among many parameters, investors and asset managers observe two of the most relevant indicators across the risky bond markets: country risk and risk or total risk. Total risk and country risk, which measure the stability of national economies through the ability to honor foreign debt payment, is one of the main parameters observed for the efficient allocation of resources in the international market. Emerging markets have proven to be increasingly efficient in terms of information, which should reveal associations between these parameters in the long run. Although the country risk can be represented by the Emerging Markets Bond Index (EMBI+), it does not present any problems regarding its determination since the prices of these securities can be observed directly in the market for these securities.

In contrast, sovereign risk refers to the credit risk of transactions that involve the state and is measured by credit rating or rating agencies such as Moody's, Fitch and Standard & Poor's. To mention those of more relevance, country risk refers to the excess return required to invest in a particular country, a risk premium, or margin which the EMBI+ indices can observe quotes for each country or by the CDS - Credit Default Swap. Resource managers observe these indicators for resource allocation. When the risk of a given market increases, investors seek to reduce the total risk of their portfolios by reallocating assets, transferring resources from bonds linked to emerging markets to developed markets in a movement called the flight to quality.

Many studies and researches have stimulated academic interest in lines of research that involve sovereign risk and country risk, especially research related to the behavior of the EMBI+ and CDS' indexes in emerging countries. Some studies seek to relate these indices to macroeconomic indicators using various methodologies, especially those that involve statistical inference. The studies that can be mentioned are Kocsis (2013) which uses principal components analysis to decompose market indicators as the EMBI+ and CDS of a representative sample of national economies; Herrera et al. (2013), which links EMBI+ with other systematic risk factors in the Mexican economy; Cristo & Gómez-Puig (2014) which presents evidence of the relationship between EMBI+ and macroeconomic variables from seven Latin American countries using autoregressive vector models with error correction mechanism and Alfaro, Medel, and Moreno (2017) that studies the country risk behavior and other three variables involving volatility and exchange rate of Latin American countries, through an autoregressive vector model. It should be noted that Alfaro et al. (2017) carry out a particular study for the Chilean economy, including the copper price in the international market, and point out that Alfaro et al. (2017) emphasize that: among Latin American countries, EMBI+ is more relevant to Brazil than to other Latin American countries.

Methods

In the construction and estimation of stochastic models necessary to achieve this work objectively, it is crucial to verify some assumptions. This way the normality is a fundamental assumption. Thus, the normality hypothesis test suggested by Jarque and Bera as described in Gujarati and Porter (2011) or Wooldridge (2012) was used. Another fundamental assumption for this research refers to stationarity. According to Gujarati and Porter (2011), it must be highlighted that nonstationarity commonly causes spurious regressions even in large samples. The stationarity test used here was the Dickey-Fuller Augmented (ADF). It is crucial to mention that the absence of stationarity can be corrected through differentiation processes. Further details can be seen in Gujarati and Porter (2011). Another fundamental statistical test for doing this work refers to the cointegration hypotheses. According to the description presented in Enders (2010), the Johansen test was selected to use among the cointegration tests available in the literature because of its suitability for estimating the autoregressive vector model shown below. Besides that, other statistical hypotheses tests inherent by the estimated models were performed.

For the conception of this research, the time series of country risk and idiosyncratic risk of the Brazilian economy was initially elaborated. The country risk is represented by its leading indicator EMBI+. It can be observed in the market, while the idiosyncratic risk determination is estimated using the market model, a stochastic linear model that relates the return of an asset to the return of the market portfolio.

The concept and determination of total risk and its components, systematic risk, and idiosyncratic risk, were established based on the Single Index Model, or market model, proposed by Sharpe (1963) that relates the returns of an asset portfolio and market portfolio returns that were previously mentioned. Within this scope of work, the risk asset portfolio is characterized by the profitability index of the Brazilian stock market, represented by the MSCI Brazil index. In contrast, the global stock market portfolio is characterized by the profitability index of the global stock market, represented by the MSCI ACWorld stock index. These two indices are calculated by the financial services company Morgan Stanley Capital Internacional (MSCI). They will be discussed below in the section dealing with the data used in this work. The market model can be described as follows:

$$R_t = \alpha + \beta R_{Mt} + e_t \quad (1)$$

where: R_t = return of the MSCI Brazil index in period t; and R_{Mt} = return on the ACWorld index or global stock market portfolio in period t. From the market model, the conditional average and conditional variance of the returns on financial assets that can be determined as follows

mean: $E(R_t|R_{Mt}) = \alpha + \beta R_{Mt}$ (2)

variance: $V(R_t|R_{Mt}) = \beta^2 V(R_{Mt}) + V(e_t)$ (3)

The total risk, measured by the conditional variance, can be divided into market risk and idiosyncratic risk according to the expression [3] above. The first installment represents a market risk, while the second installment represents an idiosyncratic risk. The beta coefficient is the main systematic risk indicator, while the market model stochastic terms variance determines the idiosyncratic risk. The estimation of the beta coefficient takes place through econometric methods applied to linear regression models, such as in the works by Scholes and Williams (1977) using a univariate model and by (A. A. Salles, 2006) using a multivariate model. The estimation of idiosyncratic risk can be done through stochastic volatility models such as the ARCH family models, particularly the GARCH suggested by Engle (1982) and Bollerslev (1986), respectively. The GARCH model, proposed by Bollerslev (1986), seeks to assimilate a standard behavior in the return series of financial assets. High values also follow large values in the following periods, not necessarily in the same direction, following a predictable process. Once alpha and beta parameters are more significant than zero, a GARCH model (p, q) can be described in its general form by the expression:

$$\sigma_t^2 = \alpha_0 + \sum_{j=1}^{q} \alpha e_{t-j}^2 + \sum_{j=1}^{p} \beta \, \sigma_{t-j}^2 \quad (4)$$

Hence, in addition to the ARCH and GARCH models, other well-known ARCH family models such as IGARCH, EGARCH, and TGARCH were tested. The model selection criterion suggested by Akaike, the AIC, which can be seen in Gujarati and Porter (2011)or Wooldridge (2012), was used to select the ARCH family model. The model selected and used in this work to estimate the idiosyncratic risk was a heteroscedastic regression model, where the stochastic terms follow a GARCH (1, 1) model, that can be described as follows:

$$R_t = \alpha + \beta R_{Mt} + e_t \, , e_t \sim Student \; (0; \sigma_t^2; \nu) \quad (5)$$

$$\sigma_t^2 = \alpha_0 + \alpha e_{t-1}^2 + \beta \sigma_{t-1}^2 \quad (6)$$

The estimation of the market model for the Brazilian market described above in [5] used the sample of returns from profitability indexes of the MSCI Brazil and MSCI ACWorld markets. Therefore, the idiosyncratic risk time series were obtained from this GARCH (1, 1) model results.

To achieve the purpose of this work, multivariate stochastic models were utilized in particular bivariate models developed from the autoregressive vector models. Presented in the econometric literature by Sims (1980), these models do not distinguish endogenous and exogenous variables and allow for the study of the relationship between two or more stochastic variables, innovations, or shocks that one variable can transmit to another or other variables and verify the short and long-run relationship between the variables involved. Also, one can verify the hypothesis of Granger causality among the variables involved (see Granger (1969) and Sims (1972)). Considering two variables of interest for this research, the idiosyncratic risk, from now on IDR, and the EMBI+ the VAR model can be described in its simplest form by the following system of equations:

$$Y_t = \beta_1 + \beta_2 Y_{t-1} + \beta_3 Z_{t-1} + \varepsilon_{1t} \quad (7)$$

$$Z_t = \beta_4 + \beta_5 Z_{t-1} + \beta_6 Y_{t-1} + \varepsilon_{2t} \tag{8}$$

where the variables Y_t and Z_t are stationary, and $\mathcal{E}_{1t}$ and $\mathcal{E}_{2t}$ have an expected value equal to zero and are orthogonal. In this work, the variable Y_t represents the variable IDR while the variable Z_t the EMBI+ Brazil. If the cointegrating hypothesis between these variable's time series is not rejected, the VAR model must be modified to VEC model or VECM. The VECM takes into consideration the cointegration between these variables through the error correction mechanism (ECM), a linear combination between Y_t and Z_t (see Salles and Almeida (2017)). This VECM model can be described in its simplest form as follow:

$$Y_t = \beta_1 + \beta_2 ECM + \beta_3 Y_{t-1} + \beta_4 Z_{t-1} + \varepsilon_{1t} \tag{9}$$
$$Z_t = \beta_5 + \beta_6 ECM + \beta_7 Z_{t-1} + \beta_8 Y_{t-1} + \varepsilon_{2t} \tag{10}$$

Results and Discussion

The primary data that formed the sample used in this work were collected on the Ipeadata website, www.ipeadata.gov.br, for EMBI +, and on the Investing.com website for the profitability indices of the Brazilian stock Market -- the MSCI Brazil Index -- and profitability of the global stock market -- MSCI All-Country World Equity Index -- or simply MSCI ACWorld Index, calculated by the financial services company Morgan Stanley Capital International (MSCI). These collected data were transformed into close weekly data, that is, the latter trading day of the week covering the period from June 7, 2009, to March 22, 2020, generating 564 weekly quotations of the equity market indices collected.

The IDR time series was obtained by estimating the market model given by the expressions shown in [5] with the stochastic terms adjusted to a Student t distribution with approximately 7 degrees of freedom. The mean, variance, and performance metrics of the estimates are shown respectively in Table 1.

Table 1. The Brazilian IDR Model Estimation Results

Mean Equation			
$MSCI\ Br_t = -0.0031 + 1.3442\ ACWORD$			
se (0.0011) (0.0451)			
Variance Equation			
$\sigma_t^2 = 0.0003 + 0.0838\, e_{t-1}^2 + 0.8927\, \sigma_{t-1}^2$			
se (0.0001) (0.0283) (0.0312)			
R-Squared	= 0.4629	Std Error of Regression	= 0.0313
Durbin-Watson Stat	= 1.8679	Sum Squared Resid	= 0.5511
Student t Degree of Freedom	= 6.71	Akaike Criterion	= - 4.2543

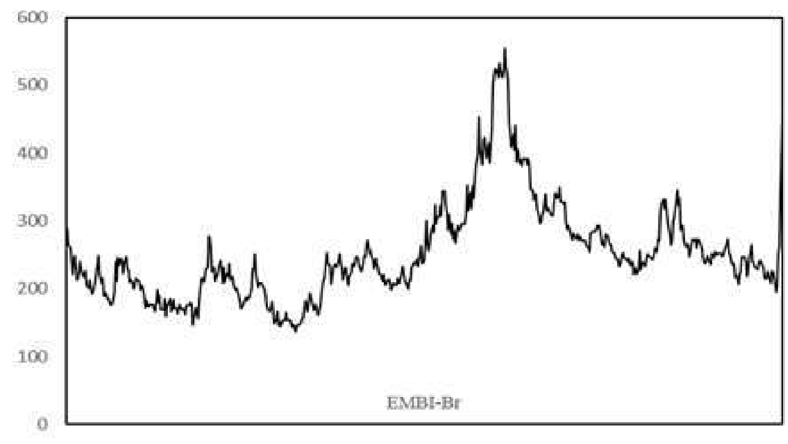

Figure 1. The Brazilian Country Risk

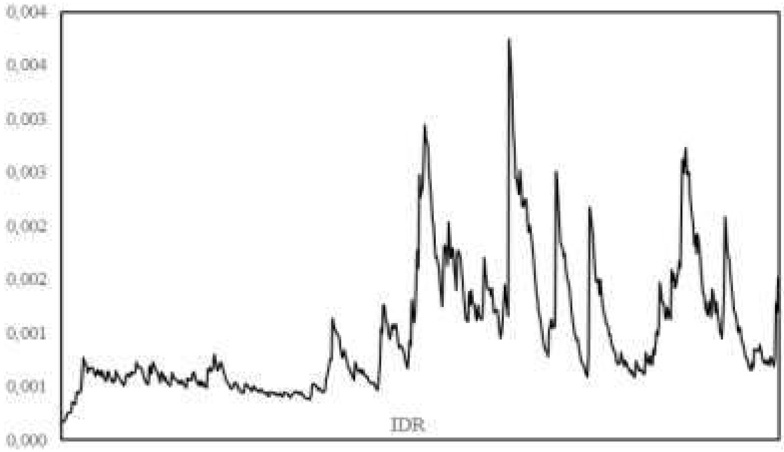

Figure 2. The Brazilian Idiosyncratic Risk

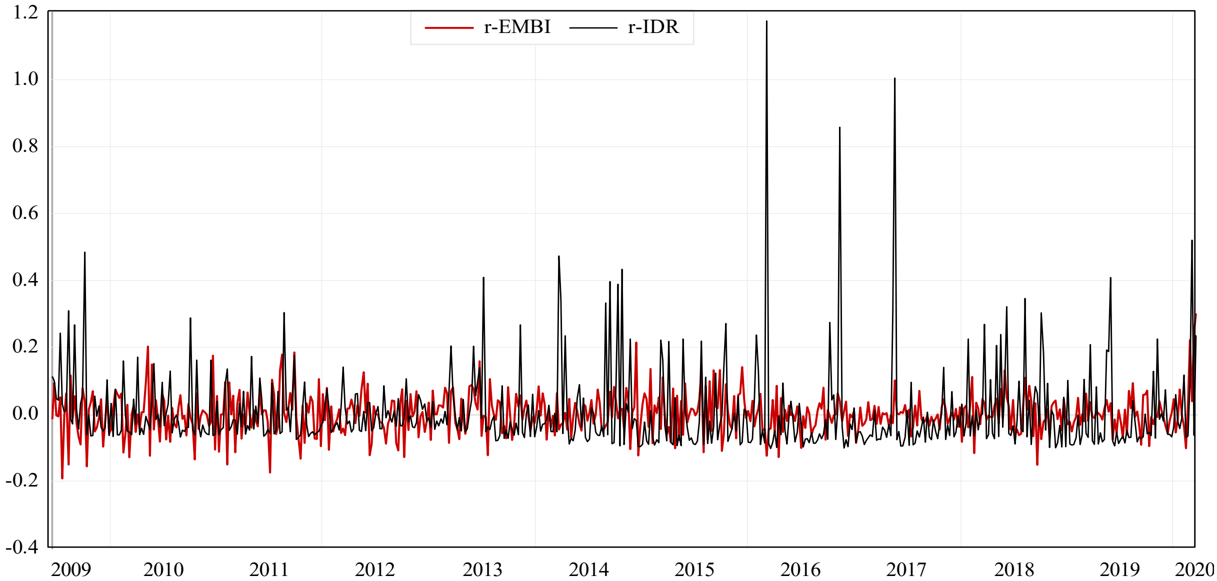

Figure 3. The Brazilian Indicator of the Country Risk and Idiosyncratic Risk Variations

From the EMBI+ and IDR time series estimates obtained, the VECM model for the elaboration of this work was implemented. The behavior and evolution of these time series during the studied period can be observed in the plots presented in Figures 1 and 2. Figure 3 presented the plots of the evolution of the variation of these indicators in the period, showing their interactions has some limitations.

Table 2. Statistical Summary of Time Series

Statistics	EMBI+ Quotes	IDR Estimates	EMBI+ Variations	IDR Variations
Mean	251.9751	0.0010	0.0009	0.0042
Median	239.0000	0.0008	0.0000	-0.0379
Maximum	554.0000	0.0037	0.3018	1.1756
Minimum	137.0000	0.0002	-0.1923	-0.1022
Std Deviation	72.5232	0.0006	0.0616	0.1265
Skewness	1.4714	1.4073	0.4388	3.9026
Kurtosis	5.9014	4.8734	4.8639	27.1381
Jarque-Bera test	400.6215	268.1563	99.5605	15096.96
(p value)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
ADF test	-2.7425	-4.4010	-16.0360	-23.2816
(p value)	(0.2198)	(0.0023)	(0.0000)	(0.0000)

Table 2 presents a time series statistical summary used to obtain the research objective, that is, to verify the iteration between the country risk and idiosyncratic risk of the Brazilian economy. To estimate the autoregressive vector model, 563 observations of EMBI+ and IDR variations were used. Table 2 presents two columns on the left showing a significant difference between the measures listed in these columns. The volatility between them represents almost 50%. The normality hypothesis of these time series presents similar and distant measures of asymmetry and kurtosis coefficients than normal distribution coefficients which are confirmed by the Jarque-Bera test, which does not allow acceptance of the normality hypothesis for these two-time series. The stationarity hypothesis for the EMBI+ time series is not accepted, while for the IDR time series, this hypothesis is not rejected. The time series summarized in Table 2 are in the last two columns on the right. These variations time series were used in the autoregressive vector model implemented in this work. It is possible to observe the different measures between these two columns, in particular in volatility time series, which shows that the variation of the IDR is greater than that of the EMBI+. The two variation time series of the two indicators do not allow rejection of the nonnormality hypothesis. Also, the hypothesis of stationarity can not be rejected for the time series of the variations or returns of these indicators, that is, the Brazilian country risk and its idiosyncratic risk. Stationarity is a fundamental assumption for the implementation of autoregressive vector models.

Table 3. The VECM Model Estimation Results

Parameter	Estimates	Std Error	Stat t	P value
β_1	-0.4843	0.0615	-7.8762	0.0000
β_2	-0.4429	0.0581	-7.6306	0.0000
β_3	-0.1547	0.0436	-3.5514	0.0004
β_4	-0.1906	0.0269	-7.0892	0.0000
β_5	-0.1016	0.0208	-4.8931	0.0000
β_6	1.3001	0.1243	10.4575	0.0000
β_7	-0.8218	0.1174	-7.0028	0.0000
β_8	-0.3801	0.0881	-4.3156	0.0000
β_9	-0.2393	0.0544	-4.4013	0.0000
β_{10}	-0.1202	0.0420	-2.8638	0.0043
Determinant Residual Covariance (DRC)=	7.41e-05			
Equation 1				
$rEMBI_t = \beta_1\ (rEMBI_{t-1} - 0{,}4942\ rIDR_{t-1}) + \beta_2\ rEMBI_{t-1} + \beta_3\ rEMBI_{t-2} + \beta_4\ rIDR_{t-1} + \beta_5\ rIDR_{t-2}$				
R-Squared = 0.4825		Mean Dependent Variable = 0.0006		
Adjusted R-Squared = 0.4787		Std Error of Dependent Variable = 0.0917		
Std Error of Regression = 0.0662		Sum Squared Resid = 2.4338		
Durbin-Watson Stat = 2.0874		Observations = 560		
Equation 2				
$rIDRt = \beta_6\ (rEMBI_{t-1} - 0{,}4942\ rIDR_{t-1}) + \beta_7\ rEMBI_{t-1} + \beta_8\ rEMBI_{t-2} + \beta_9\ rIDR_{t-1} + \beta_{10}\ rIDR_{t-2}$				
R-Squared = 0.4368		Mean Dependent Variable = 0.0003		
Adjusted R-Squared = 0.4327		Std Error of Dependent Variable = 0.1778		
Std Error of Regression = 0.1339		Sum Squared Resid = 9.9481		
Durbin-Watson Stat = 2.0829		Observations = 560		

The autoregressive vectorial models, impulse response functions, and variance decomposition were estimated to verify the short and long-run relationship between the EMBI+ and IDR variation time series. Initially, the cointegration test was performed to establish which type of autoregressive vectorial model must be estimated. Depending on the cointegration test results, the appropriate autoregressive vectorial model could be VAR or VECM. Thus, the Johansen cointegration test was performed using the time series variations of the two indicators in question. The results indicate that neither the trace statistics nor the eigenvalue statistics can reject the cointegration hypothesis between EMBI+ and IDR variation time series. The EMBI+ and IDR variations are hereafter referred to as rEMBI and rIDR. The cointegration function between these two variables, confirming the existence of a long-term association between rEMBI and rIDR. The Johansen cointegration test results indicate that the VECM model, or VAR model with error correction mechanisms, is the appropriate model to obtain necessary inferences that concern these variables' interaction.

This way, the VECM model estimation proceeded. The two equations of the VECM model estimates are shown in Table 3 in which the parameter estimated with standard errors, t statistics, and p values were presented. These results allow inferring that all parameter estimates are statistically significant and the model was estimated satisfactorily, which is confirmed with the Residual Covariance Determinant (DRC) close to zero and an AIC close to -3.79. The other model adjustment metrics and each of the autoregressive vector models with error correction can also be seen in Table 3.

From these results, tests of significance of the error correction mechanisms coefficients β_1 and β_6 were implemented and are shown in Table 2, which, with the non-rejection of the statistical significance hypothesis, indicate the existence of a long-run relationship between rEMBI + and rIDR. The Wald test the null hypothesis of these coefficients was not accepted as equal to zero, which confirms the long-run relationship between the variables rEMBI + and rIDR.

For the short-run relationship, the significance of the coefficients indicates that this hypothesis can not be rejected. Another important inference concerns the Granger causality test. The Granger causality test hypothesis points out the no rejection of bidirectional causality between the two indicators which is confirmed by the Wald test of exogeneity. There fore, it can be inferred that there is interaction in the short and long run between the two variables.

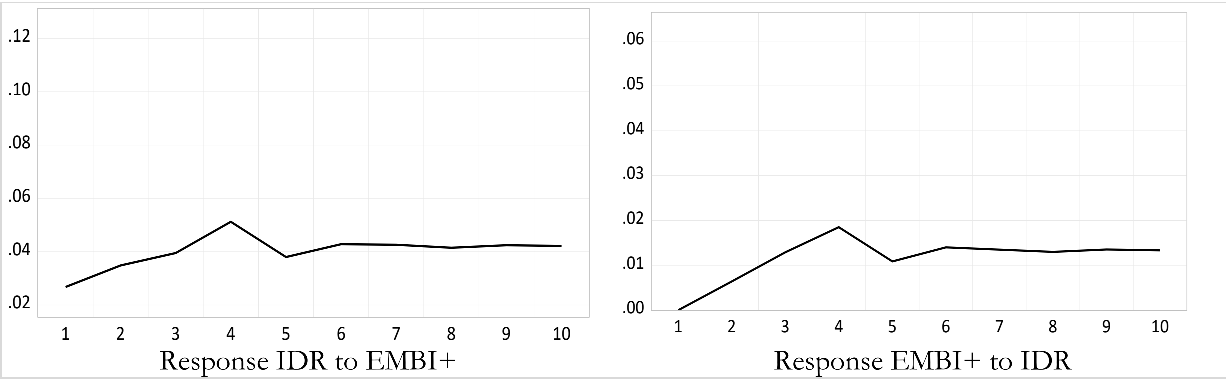

Figure 4. Impulse Response Function

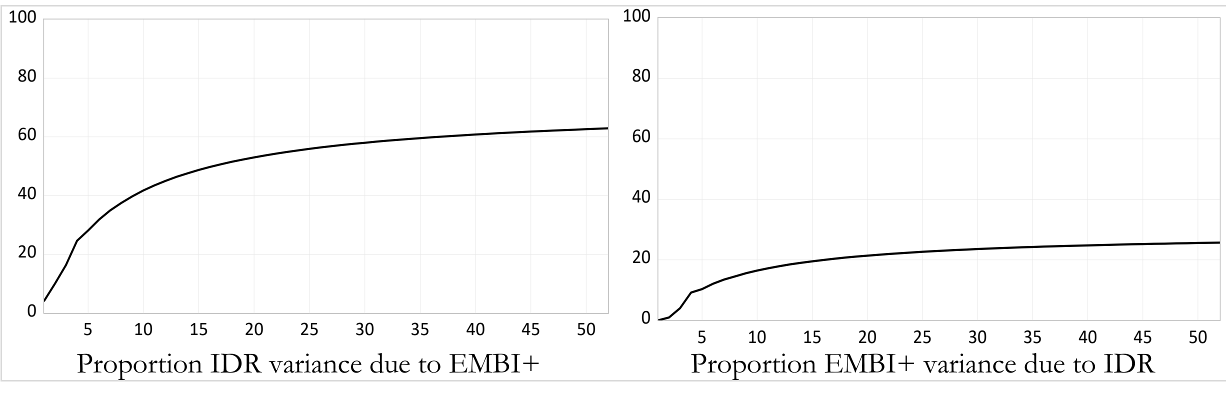

Figure 5. Variance Decomposition

Regarding the impulse response function between the variables, the plots indicated in Figure 4 allow observing how and when the response happens or lag one variable shocks on the other. That is, the responses of the variation of the country risk in the variation of the idiosyncratic risk and the variation of the idiosyncratic risk in the variation of the country risk. Concerning the variance decomposition of a variable into the other, the plot is presented in Figure 5. It can be observed that the most extensive participation in the variable variances occurs with a lag close to 10 periods or weeks, both for EMBI+ and for IDR. As Brooks (2002) observes, variance decompositions "give the proportion of movements in the dependent variables that are due to their own shocks, versus to the other variables." Additionally, Table 4 allows observing in each period listed the proportion of the explanation of the variation or of the shocks of the idiosyncratic risk due to the EMBI+ shocks and the proportion of the variation of the EMBI + due to the shocks of the variation of the idiosyncratic risk. Table 4 presents the decomposition of variance for IDR and EMBI+ time series.

The first four columns of the table show the variance decomposition of the idiosyncratic risk, namely, the first column shows the period analyzed. The second column shows the standard error of the forecast of the IDR, the third column the proportion of the variance due to shocks the IDR variable itself. In contrast, the fourth column exhibits the proportion of the variation in the explained IDR or due to EMBI+ shocks. This description could be repeated for the last four columns of Table 4 that refer to the variance decomposition of the EMBI+. It can be observed that after four weeks, approximately 25% of the variation of the EMBI+ explains the variation of the IDR while 9% IDR explains the variation of the EMBI+. After 52 weeks, approximately 62% of the variation of the EMBI+ explains the variation of the IDR, while approximately 26% IDR explains the variation in EMBI+.

Table 4. Variance Decomposition -- IDR and EMBI+

Period	Std Error	IDR EMBI+	IDR	Period*	Std Error*	EMBI+*EMBI+*	IDR*
1	0.1339	3.9973	96.0027	1	0.0662	100.0000	0.0000
2	0.1392	9.9604	90.0396	2	0.0668	99.0867	0.9133
3	0.1462	16.3381	83.6619	3	0.0720	96.0088	3.9912
4	0.1576	24.6203	75.3797	4	0.7740	90.8367	9.1633
8	0.1857	37.4914	62.5086	8	0.0913	85.4492	14.5508
12	0.2110	45.0191	54.9809	12	0.1039	82.1179	17.8821
24	0.2730	55.4279	44.5721	24	0.1346	77.5617	22.4383
36	0.3234	59.8583	40.1417	36	0.1596	75.6295	24.3706
52	0.3803	62.9030	37.0971	52	0.1877	74.3042	25.6958

*Variance Decomposition of EMBI+

The study results cannot be generalized because different developing countries and/or country groups have different economic features. Furthermore, the model was specified to test the links between only two variables. Therefore, introducing more growth factors may present different results.

Conclusion

This work aimed to estimate a proxy for a time series of the Brazilian economy idiosyncratic risk and compare the time series results obtained from the country risk measure EMBI+, which is most widespread among resource managers in the international market.

The results show similarities or associations between the two indicators with interactions in the short and long run. This way, the IDR proves to be a relevant indicator regarding the risk of economic activities implemented within the scope of the Brazilian economy and can be useful in evaluating investments in related projects. This indicator provides resource managers an alternative for obtaining estimates of the expected return required by economic agents in financing and investing in productive and infrastructure projects necessary for developing the Brazilian economy that provides greater employability and acceptable social welfare. Thus the objectives of this work were achieved.

Future research works on this theme, it is important to verify what happens in other economies, which refer to the idiosyncratic risk behavior and the associations and interactions of country risk indicator and idiosyncratic risk and their behavior in economies in general. Furthermore, it is meant to highlight the use of other methodological approaches.

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Commercial banks regulation and intermediation function in an emerging market

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Abstract

Purpose — This paper investigates the effect of commercial bank regulations, namely the price, product, and geographic regulations, on the intermediation function of commercial banks in Nigeria.

Methods — Using secondary data from 1986 to 2017 from the Central Bank of Nigeria (CBN) and the World Bank, this study employs the Autoregressive Distributive Lag (ARDL) model and Granger causality framework.

Findings — This paper provides evidence of a long-run relationship between commercial bank regulation and intermediation function represented by private sector credit to RGDP (regional gross domestic product). It also finds that commercial banks' regulation index through price, product, and geographic regulation has a positive relationship with intermediation function. Furthermore, the long-run relationship between commercial bank regulation and intermediation function described by private sector credit to RGDP is affirmed.

Implication — The Central Bank of Nigeria (CBN) needs to relax the product regulation to allow commercial banks to engage in various conventionally non-banking activities.

Originality — The paper contributes to the literature by ascertaining the commercial banks' intermediation function to Nigeria's economic growth and development.

Keywords — Commercial bank regulation, geographic regulations, intermediation function.

Introduction

No economy will tear the part of growth and development without the intermediation function of financial institutions. Realistically, coupled with the uncertainty that characterized the macroeconomic environment of emerging economies, to presume there would be sustainability in growth in the economy shorn of the intermediation function of financial institutions is to assume our difficulties and worries away. Financial institutions, within the framework of the financial markets and instrumentality of the financial assets, intermediate in funds to bring 'together' the surplus and deficit economic agents in such a manner as to resolve existing financial resources' imbalance among them (Ezirim & Moughalu, 2002). Credit allocation to different sectors of the economy leans on the operation and structure of the financial institutions. Hence productive economic activities rely virtually on the intermediation process of financial institutions, especially

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for emerging economies where the financial markets are in their developing stages, and firms vehemently trust the banking system for external finances. This declaration cannot be put to contention, for within the realm of theoretical literature, Nwakoby and Ananwude (2016) explicitly stated that when financial system and instruments of operation are strengthened, transaction and information costs in the economy will decline tremendously, which in turn, influence savings rate, investment decision, and innovative technological ventures. Strengthening financial institutions have roots in the legal and regulatory fundamentals of the economy. Sound legal and regulatory environments lower financial intermediation costs, whereas poor and weak legal and regulatory environments, result in intensification in financial intermediation costs, an indispensable reticence for financial deepening in the economy.

Empirical studies on the nexus between commercial banks regulation and intermediation function of financial institutions are relatively few for developed economies (see Bottazzi, Da Rin, & Hellmann, 2009; Kale, Eken, & Selimler, 2015; Vittas, 1991), but there abound large theoretical and documentary pieces of evidence. Concerning emerging economies, empirical studies are also scarce due to scholars' reluctance to research this area attributable to the immeasurability of proxies. In the course of review of empirical studies in the Nigerian setting, it was observed with dismay that there is a shortage of studies on the effect of banking regulation on the intermediation function of commercial banks. The only online available studies in this area at the time this study was undertaken were Ezirim and Moughalu (2002) and Ezirim, Muoghalu, and Emenyonu (2004). They established that commercial bank regulation exerts a significant effect on their intermediation function. By time or period implication, there is a need for further study in this area which this study seeks to fulfil. This study is slightly distinguished from Ezirim and Moughalu (2002) in three aspects. First, it specifically centered on commercial bank regulation, whereas Ezirim and Moughalu (2002) studied both the legal and regulatory environments. Secondly, it covered only the period of deregulation (1986 to 2017) compared to the period of the regulation (1970 to 1985) and deregulation (1986 to 2000) undertaken by Ezirim and Moughalu (2002). Finally, it introduced credit to the private sector relative to real gross domestic product to measured financial intermediation and total assets of commercial banks concerning the real gross domestic product.

Regulation plays a critical role in the efficient intermediation function of commercial banks. In its simplest form, regulation is a set of rules that guides the operation of commercial banks intending to ensure viability, sustainability, and stability of the financial system in such a manner that commands public confidence in the payment system. Regulation or supervision has to do with the laid down rules and regulations that control the operation of the financial institutions. Owing to the expanded competition, the borders between commercial banks are blurring, financial innovations are duplicating off-balance-sheet exercises, and internationalization is rendering control by financial specialists increasingly troublesome (Heremans & Pacces, 2011), hence the need for regulation of commercial banks. The rationale for government regulation of the financial system, which differs from country to country, is often tied to the disastrous effect of market failure and imperfection in market competition practices in the financial sector. The type of regulation imposed by the government influences financial institutions' type of financial products and services. Regulation in terms of interest rate affects the overall cost of capital in the economy. Regulation of permissible and non-permissible activities eminently affects the financial institutions' income, liquidity, and solvency position. Reserve requirements also alter the magnitude of commercial credit banks extends to the economy. Gorton and Winton (2002) avowed that government regulation by the provision of deposit insurance and intervention into banking markets, including bank supervision and examination, limitations on bank activities, capital requirements, charter requirements and entry restrictions, closure rules, and other rules are motivated by the fact that financial institutions, particularly the banks are regarded as flawed institutions prone to harmful banking panics. From theoretical and empirical shreds of evidence, the benefits of government intervention in the operation of commercial banks through regulation and supervision are indisputable.

Financial intermediation may be interpreted in different ways by scholars but for all point to the proficient mobilization and allocation of scarce economic resources. Financial intermediation, as performed by commercial banks, which varies from country to country depending on the level of development in the financial system, is the process of bringing together surplus spending units (savers) and deficit spending units (borrowers) in an economy (Kwakye, 2012). From Ezirim (1999), financial intermediation is the process whereby financial institutions creating financial assets within the framework of the financial markets, bring together the surplus economic units and the deficit economic units to resolve the financial imbalance employing a price-related compensation mechanism referred to as interest rate. In Nigeria, commercial banks' intermediation function is largely performed compared to the financial markets and non-bank financial institutions operating in the country.

The theory of financial intermediation, in a nutshell, states that the effective and efficient intermediation function of financial institutions provides the economy with the needed funds for productive activities, which lead to growth and development. The theory of financial intermediation is seen as the finance-led growth theory: supply leading hypothesis advanced by Schumpeter in 1911 and made popular by Shaw (1973) and McKinnon (1973). Informational asymmetry and agency theory is considered the theoretical foundation of the theory of financial intermediation. The modern theory of financial intermediation is conceptualized that financial intermediaries serve to decrease costs of transaction and instructive asymmetries, for as improvements in data innovation, deregulation, development of financial markets, etc. tend to decrease transaction costs and information asymmetries, the financial intermediation hypothesis might conclude that intermediation gets to be futile (Scholtens & Wensveen, 2003). In this study, Ferreira de Freitas (2014) and Nwakoby and Ananwude (2016) instigated the discussion of the theory of financial intermediation using two concepts: information asymmetry and transaction cost.

Information asymmetry entails a situation where lenders of funds have inadequate information about the firms requiring the funds or the projects they utilize. There is no doubt that in the present-day business environment, especially in emerging economies, Nigeria, for instance, borrowers tend to have more information on the risk related to the projects they commit to lenders' funds compared to the savers themselves. The disclosure from Nwakoby and Ananwude (2016) affirmed that information asymmetry gives rise to the ethical risk and antagonistic determination problems, driving the plausibility of reducing the capability of efficiency of allocation of funds deficit units from excess units.

Transaction cost dwells on the cost of accessing financial products and services. Lenders would want to be paid interest for giving out their surplus funds. Likewise, the financial intermediaries would want income for lending mobilized funds to the ultimate users: borrowers. The interest rate required by depositors and the fee charged by the financial intermediaries through enforcement, monitoring, verification, and search costs, etc., determines the overall transaction cost in an economy. As said by Nwakoby and Ananwude (2016), the transaction cost approach is a follow up the perfect market condition where according to the neoclassical economists, prices in the market cannot be influenced by one partaker, conditions for lending and borrowing for all partakers are indistinguishable. All information concerning factors and components capable of influencing the present or future values of financial securities are instantaneously at the disposal of all partakers.

Kale et al. (2015) evaluated the impacts of regulations, macroeconomic changes, and political occasions on the effectiveness of the Turkish banks amid the period 1997-2013 when pivotal changes were experienced. A two-stage strategy was utilized. To begin with, the efficiency changes of each bank and the entire segment were measured by a DEA-based Malmquist Productivity Index (DEA-MPI). In common, a modern macroeconomic environment, especially new regulations, has positive impacts on efficiency. More tightly regulations, checking, confinements, solid supervision, more capital, and new reforms positively affect effectiveness.

Berka and Zimmermann (2018) assessed the impact of the Basel Accord on financial intermediation. They found that a monetary policy approach increments credit volume indeed when the economy is in great shape. In contrast, an active capital necessity approach is also viable if it suggests fixing of regulation in bad times.

Zheng, Rahman, Begum, and Ashraf (2017) assessed the impact of capital necessities on the cost of financial intermediation and bank profitability employing a board dataset of 32

Bangladeshi banks over 2000 to 2015. By utilizing the dynamic panel Generalized Method of Moments (GMM) estimator, the study supported that higher bank regulatory capital ratios decrease the costs of financial intermediation and increase bank profitability. Moreover, they found that switching from BASEL I to BASEL II has no quantifiable effect on the costs of financial intermediation and bank profitability in Bangladesh.

Antunes and Moraes (2017) investigated the behavior of financial intermediation through the analysis of a panel of 101 Brazilian banks. The results indicated an increase in non-performing loans and a tight monetary policy increase financial friction and reduced financial intermediation.

Bonner and Eijffinger (2016) examined the impact of liquidity regulation on bank intermediation applying regression discontinuity designs. Employing a special dataset on Dutch banks, it appeared that a liquidity requirement causes long-term borrowing and loaning rates as well as requests for long-term interbred advances to extend. Lower levels of total liquidity increment the assessed impacts.

Employing a panel data set of commercial banks from eight major Asian economies over the period 2001-2010, Deng, Casu, and Ferrari (2014) investigated how the coexistence of progression and prudential regulation influences banks' cost characteristics. Discoveries appeared that liberalization of bank interest rates and increased presence of foreign banks had had a positive and critical effect on technological progress and cost-efficiency. Results too uncovered that prudential regulation might unfavorably influence bank cost performance.

Ferreira de Freitas (2014) ascertained the effect of bank regulatory capital on liquidity creation for the economy based on a sample of Euro area banks for 2006-2012. The study found that higher regulatory capital negatively impacts liquidity creation. However, no evidence was found that the relationship between regulatory capital and liquidity creation differs from bank size or during a crisis period.

Wanjiru (2012) looked at the effect of financial regulation on the financial performance of deposit-taking microfinance institutions in Kenya. The research design that was used in this study was both a cross-sectional and descriptive survey method. The study concluded that the supportive deposit-taking microfinance regulations of 2008 led to the improved financial performance of deposit-taking microfinance institutions.

Demirgüç-Kunt, Laeven, and Levine (2004) looked at the impact of bank regulations, market structure, and national institutions on bank net interest margins and overhead costs, utilizing over 1,400 banks over 72 nations whereas controlling for bank-specific characteristics. The information showed that more tightly regulation on bank entry and bank activities boost financial intermediation cost.

Ezirim and Moughalu (2002) empirically ascertained the effect of the legal and regulatory environments on financial intermediation in an emerging Sub-Saharan economy, with evidence drowned from the Nigerian commercial banks. The method employed included the construction, estimation, and analysis of econometric models. Utilizing yearly time-series information (1970–2000), the findings appeared that the legal and regulatory environments exerted a noteworthy impact on the financial intermediation operations of commercial banks in Nigeria.

Park and Sehrt (2001) used Chinese provincial data from 1991 to 1997 to test whether financial reforms in the mid-1990s increased efficient intermediation by different financial institutions. The results indicated that financial intermediation in China is far from efficient and that financial reforms in the mid-1990s have not reversed the trend of worsening bank performance.

Having justified our motivation for undertaking this study and reviewed relevant literature, the next section of this article (methods) reviews the data, methodological steps, processes, and after that, result and discussion. The last section depicts the conclusion.

Methods

Firstly, we followed the step of determining the descriptive properties of the data employed in the study. That notwithstanding, we went further to ascertain the correlation matrix between commercial bank regulation and intermediation function variables.

In the second step, we checked for the stationarity properties of the data. This is because the non-stationarity of time series data leads to spurious regression results, which cast a dent in the statistical regression output's reliability concerning econometric assumption. To this end, the study used the Augmented Dickey-Fuller (ADF) Test and Phillips Perron (PP) tests of a unit root.

Thirdly, we estimated the long-run relationship between commercial bank regulation and intermediation function and the nature of the relationship in the long-run. The short-run relationship between the variables of interest was also evaluated. These estimations were carried out using the Autoregressive Distributive Lag (ARDL) technique. The choice of ARDL as an econometric tool hinges on the fact that it takes into account a different order of stationarity achieved by time-series data.

The fourth step led to evaluating the robustness of the model via serial correlation LM test, heteroscedasticity, and Ramsey Reset Specification test. These tests confirm the authenticity of the Autoregressive Distributive Lag (ARDL).

Finally, we employed the granger causality framework as against the Ordinary Least Square (OLS) technique to examine the effect of commercial bank regulation on intermediation function. Jeff-Anyeneh, Anawude, Ezu, and Nnoje (2020) had it that the Granger causality framework is superior to the Ordinary Least Square (OLS) technique since the Granger causality framework is more efficient and effective in determining whether a time-series variable is useful in predicting or forecasting another. The OLS ordinarily tests for the "mere" relationship between variables. Two variables may relate without one causing changes in the other, hence the vehement reliability of the granger causality framework in this regard.

The data used in the analysis were secondary and were sourced from the Central Bank of Nigeria Statistical Bulletins of various issues and the World Bank. The scope of the study spans from 1986 to 2017. The intermediation function of commercial banks was defined in terms of Private Sector Credit to Real Gross Domestic Product (CPSR) and Total Assets of Commercial Banks to Real Gross Domestic Product (TAR). The inclusion of private sector credit to the real gross domestic product as a proxy for financial intermediation is on the fact that in Nigeria, the depth or degree of intermediation function of commercial banks is seen in the magnitude of fund extended to the private sector from funds mobilized from surplus units, and was guided by the work of Nwakoby and Ananwude (2016). Commercial banks regulation indices are Prime Lending Rate (PLR) for price regulation; Liquidity Ratio (LR) for product regulation; Number of Commercial Banks Branches (NB) and Commercial Bank Density (DEN): the ratio of commercial banks branches to total population for geographic regulation.

The realization of the objective of this study followed an estimation of a linear regression equation, and a reminiscent of Ezirim and Moughalu (2002) expressed as:

$$FII = f(ILE, GI, PDI, PCI\ ) \tag{1}$$

Where: FII: Financial intermediation index; ILE: Index of the legal environment; GI: Geographic regulation index; PDI: Product regulation index; PCI: Price regulation index

The model has been modified by removing the legal regulation index. In this regard, equation (1) is modified as:

$$FI = f(PCR, PDR, GR\ ) \tag{2}$$

The linear and log-linear form of equation (1), having inputted the various measurement of regulation, are stated as:

For linear function:

$$CPSR_t = a_0 + a_1PLR_t + a_2LR_t + a_3NB_t + a_4DEN_t + \varepsilon_{1t} \tag{3}$$
$$TAR_t = a_0 + a_1PLR_t + a_2LR_t + a_3NB_t + a_4DEN_t + \varepsilon_{1t} \tag{4}$$

For log-linear function:

$$LogCPSR_t = \beta_0 + \beta_1LogPLR_t + \beta_2LogLR_t + \beta_3LogNB_t + \beta_4LogDEN_t + \varepsilon_{2t} \tag{5}$$
$$LogTAR_t\ = \beta_0 + \beta_1LogPLR_t + \beta_2LogLR_t + \beta_3LogNB_t + \beta_4LogDEN_t + \varepsilon_{2t} \tag{6}$$

Where:
$CPSR$: Private sector Credit to RGDP; TAR: Total assets of commercial banks to RGDP; PLR: Prime lending rate; LR: Liquidity ratio; NB: Number of branches of commercial banks; and DEN: Density of commercial banks.

Parameters and elasticities of the models are described by a_1, β_0,$<_0/>_0$; a_2, β_2, $<_0$, a_3, β_3,$<_0$, a_4, β_4 $<_0$ and a_0 and β_0 respectively, while error terms are defined by ε_{1t} and ε_{2t}.
The application of the ARDL method of estimation would mean that the models will be partially adjusted in line with the Partial Adjustment Model (PAM). Hence partial adjustment model for equation 3 – 6 are written is:
For linear function:

$$CPSR_t = \varphi_0 + \varphi_1 PLR_t + \varphi LR_t + \varphi_3 NB_t + \varphi_4 DEN_t + \varphi_5 CPSR_{t-1} + \varepsilon_{1t} \quad (3a)$$

$$TAR_t = \varphi_0 + \varphi_1 PLR_t + \varphi LR_t + \varphi_3 NB_t + \varphi_4 DEN_t + \varphi_5 TAR_{t-1} + \varepsilon_{1t} \quad (4a)$$

For log-linear function:

$$LogCPSR_t = \omega_0 + \omega_1 LogPLR_t + \omega_2 LogLR_t + \omega_3 LogNB_t + \omega_4 LogDEN_t + \omega_5 LogCPSR_{t-1} + \varepsilon_{2t} \quad (5a)$$

$$LogTAR_t = \omega_0 + \omega_1 LogPLR_t + \omega_2 LogLR_t + \omega_3 LogNB_t + \omega_4 LogDEN_t + \omega_5 LogTAR_{t-1} + \varepsilon_{2t} \quad (6a)$$

Where:
$\lambda = 1 - \varphi_5$; $m_0 = \varphi_0/\lambda$; $m_1 = \varphi_1/\lambda$; $m_2 = \varphi_2/\lambda$; $m_3 = \varphi_3/\lambda$; $m_4 = \varphi_4/\lambda$. The short-run and long-run coefficients are described in terms of φ_{1-4} and m_{1-4} respectively.
$\lambda = 1 - \omega_5$; $l_0 = \omega_0/\lambda$; $l_1 = \omega_1/\lambda$; $l_2 = \omega_2/\lambda$; $l_3 = \omega_3/\lambda$; $l_4 = \omega_4/\lambda$. The short-run and long-run coefficients are defined in terms of ω_{1-4} and l_{1-4} respectively.

Results and Discussion

The mean, median, maximum, standard deviation, skewness, kurtosis, Jarque-Bera, p-value, and the number of observations were utilized to describe the descriptive properties of the data, as shown in Table 1. Commercial banks regulation index concerning prime lending rate, liquidity ratio, number of commercial banks branches, and density reveal the mean of 18.79, 45.29, 3402.59, and 0.0023, while intermediation function defined in term of private credit to RGDP and total assets of commercial banks to RGDP has 11.85 and 158.23 respectively. The median of the data was defined in terms of 8.25, 92.17, 17.96, 44.65, 2703.50, and 0.002, while standard deviation as 6.02, 167.25, 3.78, 8.84, 1556.93, and 0.0006 accordingly for CPSR, TAR, PLR, LR, NB, and DEN. The maximum and minimum were explained by 23.10 and 6.20 for CPSR, 486.24 and 2.61 for TAR, 29.80 and 10.50 for PLR, 64.10 and 29.10 for LR, 5805 and 1367 for NB, and 0.034 and 0.001 for DEN. The data were positively skewed toward normality but from the Kurtosis statistic, only PLR was found to be leptokurtic. The p-values of the data (significant at 5% level) have proved beyond a reasonable doubt that the data followed a normal distribution. With this, the regression output would be said to be free from any outlier that may cast a dent in the data output.

Table 1: Data Descriptive Properties

	Mean	Median	Maximum	Minimum	Std. Dev.	Skewness	Kurtosis	Jarque-Bera	P-value	Obs
CPSR	11.856	8.250	23.10000	6.200000	6.022803	0.806228	1.860801	7.197050	0.0343	32
TAR	158.23	92.175	486.2400	2.610000	167.2493	0.645843	1.788378	8.181974	0.0135	32
PLR	18.787	17.965	29.80000	10.50000	3.778648	0.947693	4.598371	8.196368	0.0166	32
LR	45.287	44.650	64.10000	29.10000	8.848579	0.213314	2.763320	7.317371	0.0432	32
NB	3402.594	2703.50	5805.000	1367.000	1556.934	0.445143	1.564329	6.805014	0.0491	32
DEN	0.002	0.002	0.003400	0.001430	0.000612	0.431968	1.742336	9.104141	0.0099	32

Source: Output data from E-views 10.0

The highest correlation (-0.30) for commercial banks regulation index was observed for price and geographic regulation: prime lending rate and the number of commercial bank branches. This is low and within the acceptable range of no multi-collinearity issue in a model. In this case, this study is convinced that there is no multi-collinearity problem between the measures of commercial banks regulation index as divulged in Table 2.

Table 2: Correlation Matrix

	CPSR	TAR	PLR	LR	NB	DEN
CPSR	1.000					
TAR	0.938	1.000				
PLR	-0.378	-0.367	1.000			
LR	-0.139	-0.017	-0.079	1.000		
NB	0.927	0.968	-0.303	-0.076	1.000	
DEN	0.900	0.907	-0.234	-0.199	0.975	1.000

Source: Output data from E-views 10.0

Table 3. ADF Test Result at Level

Variables	Intercept	Trend and Intercept	None	Remark
CPSR	-0.798 (0.80)	-2.097 (0.53)	0.461 (0.80)	Not Stationary
TAR	0.955 (0.99)	-2.033 (0.56)	1.430 (0.96)	Not Stationary
PLR	-4.569 (0.00)*	-5.886 (0.00)*	-1.357 (0.16)	Stationary
LR	-3.339 (0.02)**	-3.286 (0.08)	-0.311 (0.57)	Stationary
NB	-0.267 (0.92)	-2.687 (0.25)	2.266 (0.99)	Not Stationary
DEN	-0.994 (0.74)	-3.262 (0.09)	1.154 (0.93)	Not Stationary

Source: Output data from E-views 10.0
Note: The optimal lag for the ADF test is selected based on the Akaike Info Criteria (AIC), p-values are in parentheses where (*) and (**) denote significance at 1% and 5%, respectively.

Table 4. ADF Test Result at First Difference

Variables	Intercept	Trend and Intercept	None	Remark
CPSR	-5.846 (0.00)*	-5.817 (0.00)*	-5.783 (0.00)*	Stationary
TAR	-3.084 (0.04)**	-4.059 (0.01)*	-2.387 (0.01)*	Stationary
PLR	-4.945 (0.00)*	-4.894 (0.00)*	-4.698 (0.00)*	Stationary
LR	-6.109 (0.00)*	-5.971 (0.00)*	-6.216 (0.00)*	Stationary
NB	-4.053 (0.00)*	-3.990 (0.02)**	-3.499 (0.00)*	Stationary
DEN	-4.120 (0.00)*	-4.041 (0.02)**	-3.969 (0.00)*	Stationary

Source: Output data from E-views 10.0
Note: The optimal lag for the ADF test is selected based on the Akaike Info Criteria (AIC), p-values are in parentheses where (*) and (**) denote significance at 1% and 5% respectively.

Table 5. PP Test Result at Level

Variables	Intercept	Trend and Intercept	None	Remark
CPSR	-0.798 (0.81)	-2.097 (0.53)	0.544 (0.83)	Not Stationary
TAR	1.419 (0.99)	-1.443 (0.83)	2.743 (0.99)	Not Stationary
PLR	-4.687 (0.00)*	-5.793 (0.00)*	-0.155 (0.62)	Stationary
LR	-3.251 (0.02)**	-3.191 (0.10)	0.203 (0.74)	Stationary
NB	-0.267 (0.92)	-1.739 (0.71)	2.266 (0.99)	Not Stationary
DEN	-1.080 (0.71)	-1.794 (0.68)	1.155 (0.93)	Not Stationary

Source: Output data from E-views 10.0
Note: In determining the truncation lag for the PP test, the spectral estimation method selected is Bartlett kernel and Newey-West method for Bandwidth, p-values are in parentheses where (*) and (**) denote significance at 1% and 5%, respectively.

This study applied the Augmented Dickey-Fuller (ADF) Test and Phillips Perron (PP) tests of unit root to determine the stationarity of the data. Test for unit root was in level and first difference, and included in the test equation were: intercept, trend and intercept, and none. The data were all stationery from the unit test result in Tables 4 and 6; the data were all stationary at first difference as against non-stationarity for all data in Tables 3 and 5.

Table 6. PP Test Result at First Difference

Variables	Intercept	Trend and Intercept	None	Remark
CPSR	-5.846 (0.00)*	-5.818 (0.02)**	-5.783 (0.00)*	Stationary
TAR	-2.801 (0.05)**	-3.109 (0.04)**	-2.256 (0.03)**	Stationary
PLR	-10.197 (0.00)*	-9.867 (0.00)*	-10.400 (0.00)*	Stationary
LR	-13.200(0.00)*	-12.519 (0.00)*	-13.363 (0.00)*	Stationary
NB	-4.064 (0.00)*	-4.001 (0.02)**	-3.489 (0.00)*	Stationary
DEN	-4.125 (0.00)*	-4.046 (0.02)**	-3.964 (0.00)*	Stationary

Source: Output data from E-views 10.0

Note: In determining the truncation lag for the PP test, the spectral estimation method selected is Bartlett kernel and Newey-West method for Bandwidth, p-values are in parentheses where (*) and (**) denote significance at 1% and 5%, respectively.

If the f-statistic of the bound test is higher than the upper bound critical value at a 5% significance level, the null hypothesis of no long-run relationship is rejected. The revelation in Table 7 shows a long-run relationship between commercial bank regulation and intermediation function defined in terms of private sector credit to RGDP. Still, such is not the case for describing intermediation function by total assets of commercial banks to RGDP.

Table 7. Bound Test for Regulatory Index and Financial Intermediation

	FI by CPSR	FI by TAR
F-Statistic	3.69	2.37
Lower Bound @ 5% Critical Value Bound	2.56	2.56
Upper Bound @ 5% Critical Value Bound	3.49	3.49

Source: Output data from E-views 10.0

Table 7 provided the existence of a long-run relationship between commercial bank regulation and intermediation function represented by private sector credit to RGDP. By implication, the determination of the nature of the long-run relationship along the line the speed of adjustment becomes imperative. From the result in Table 8, commercial banks' regulation index through price, product, and geographic regulation has a positive relationship with intermediation function. For the speed of adjustment, the ECM is rightly signed and significant at a 5% level of significance. In essence, there is a tendency by the model to move towards equilibrium following disequilibrium in previous periods. Furthermore, the long-run relationship between commercial bank regulation and intermediation function described by private sector credit to RGDP is affirmed.

The relationship in the short run between commercial bank regulation and intermediation function was appraised by the ARDL estimation. The different order of integration of the data guided the choice of ARDL. Adjusted R-squared, F-statistic, and Durbin Watson statistic are three global model utility employed. Similarly, the relative statistics of the variables were not ignored. Table 9 provides evidence of an insignificant negative relationship between prime lending rate, liquidity ratio, and intermediation function defined in terms of private sector credit to RGDP. It also provides evidence of a positive relationship between the number of commercial bank branches, commercial banks density, and intermediation function. Intermediation function, when described by total assets of commercial banks to RGDP, has a negative relationship with prime lending rate and density but a positive relationship with liquidity ratio and the number of commercial banks branches.

Table 8. ARDL Error Correction CPSR→PLR, LR, NB and DEN

ECM Regression				
Variable	Coefficient	Std. Error	t-Statistic	Prob.
D(CPSR(-1))	1.019	0.246	4.139	0.009
D(CPSR(-2))	0.184	0.150	1.228	0.274
D(PLR)	0.955	0.185	5.155	0.004
D(PLR(-1))	0.417	0.129	3.245	0.023
D(PLR(-2))	0.871	0.164	5.308	0.003
D(PLR(-3))	0.559	0.111	5.037	0.004
D(LR)	0.002	0.039	0.056	0.958
D(LR(-1))	-0.050	0.032	-1.598	0.171
D(LR(-2))	-0.112	0.029	-3.810	0.013
D(NB)	0.005	0.004	1.157	0.300
D(NB(-1))	0.046	0.007	6.558	0.001
D(NB(-2))	0.007	0.006	1.121	0.313
D(NB(-3))	0.021	0.005	3.947	0.011
D(DEN)	-6522.942	6134.716	-1.063	0.336
D(DEN(-1))	-78964.410	12020.090	-6.569	0.001
D(DEN(-2))	-24916.780	10442.740	-2.386	0.063
D(DEN(-3))	-39545.820	8414.873	-4.700	0.005
CointEq(-1)*	-2.071	0.311	-6.654	0.001
Long Run Coefficient				
PLR	0.359	0.224	1.601	0.170
LR	0.007	0.067	0.112	0.915
NB	0.001	0.002	0.731	0.497
DEN	6529.154	4373.435	1.493	0.196
C	-16.794	7.093	-2.368	0.064

Source: Output data from E-views 10.0

Table 9. ARDL Regression Result of Regulatory Environment and Financial Intermediation

Variables	FI by CPSR		FI by TAR	
	Coefficient	Prob.	Coefficient	Prob.
CPSR(-1); TAR(-1)	0.491	0.003	1.238	0.000
PLR	-0.207	0.099	-0.801	0.633
LR	-0.036	0.495	0.293	0.643
NB	0.001	0.372	0.033	0.418
DEN	834.820	0.805	-25414.210	0.645
C	5.151	0.368	-16.332	0.828
Adjusted R-squared	0.895		0.985	
F-statistic	52.207		205.606	
Prob(F-statistic)	0.000		0.000	
Durbin-Watson stat	2.165		1.953	

Note: FI by CPSR and FI by TAR defined financial intermediation by the credit to private sector ratio to RGDP and total assets of commercial banks ratio to RGDP, respectively.
Source: Output data from E-views 10.0

If commercial banks regulation index: price, product, and geographic regulations are held constant, intermediation function by CPSR would be 5.15, whereas intermediation function by TAR would be -16.33. CPSR model of intermediation function would rise by a factor of 0.0014 and 834.82, respectively, owing to a unit increase in the number of commercial bank branches and commercial banks density. In contrast, the CPSR intermediation function model would decline by 20.72% and 3.58%, accordingly following a percentage rise in prime lending rate and liquidity ratio. For the TAR model of intermediation function, a unit increase in prime lending rate and density of commercial banks result in 0.80 and 25414.21-factor decrease in intermediation function. In contrast, a percentage appreciation in liquidity ratio and the number of commercial banks branches lead to 29.29% and 3.32% increase in intermediation function.

Concerning the adjusted R-square, TAR model of intermediation function (98.50%) is higher than CPSR (89.51%) model of intermediation function. Commercial banks regulation index: price, product, and geographic regulation significantly explained the changes in both CPSR and TAR intermediation function model (p-values of F* for CPSR and TAR < 0.05). There is no autocorrelation issue in both models of intermediation function (CPSR and TAR) as the Durbin Watson coefficient of 2.16 and 1.95 absolved the models of autocorrelation problem.

The robustness of the models' visa viz: serial correlation and heteroskedasticity in Tables 10 and 11 absolve the models estimated of serial correlation and heteroskedasticity problems (p-values > 0.05). The residual diagnostic in Table 12 provided evidence that the models were well-specified, and no misspecification issue was observed (p-value > 0.05).

Table 10. Serial Correlation LM Test

Estimated Equations	Obs*R-squared	F-statistic	P-value
CPSR →PLR + LR + NB + DEN	0.700	0.266	0.769
TAR → PLR + LR + NB + DEN	1.371	0.422	0.663

Source: Output data from E-views 10.0

Table 11. Heteroskedasticity Test

Estimated Equations	Obs*R-squared	F-statistic	P-value
CPSR →PLR + LR + NB + DEN	8.261	1.817	0.146
TAR → PLR + LR + NB + DEN	0.884	8.562	0.556

Source: Output data from E-views 10.0

Table 12. Ramsey Reset Specification

Estimated Equations	F-statistic	df	P-value
CPSR →PLR + LR + NB + DEN	0.216	(1, 24)	0.646
TAR → PLR + LR + NB + DEN	1.215	(1, 18)	0.285

Source: Output data from E-views 10.0

Table 13. Granger Causality Analysis

Financial Intermediation by CPSR				
Null Hypothesis:	Obs	F-Statistic	Prob.	Remarks
PLR does not Granger Cause CPSR	31	0.264	0.612	No Causality
CPSR does not Granger Cause PLR		2.587	0.119	No Causality
LR does not Granger Cause CPSR	31	0.000	0.998	Causality
CPSR does not Granger Cause LR		0.183	0.672	No Causality
NB does not Granger Cause CPSR	31	11.030	0.003	Causality
CPSR does not Granger Cause NB		0.024	0.878	No Causality
DEN does not Granger Cause CPSR	31	8.093	0.008	Causality
CPSR does not Granger Cause DEN		0.586	0.451	No Causality
Financial Intermediation by TAR				
PLR does not Granger Cause TAR	31	0.462	0.502	No Causality
TAR does not Granger Cause PLR		5.144	0.031	Causality
LR does not Granger Cause TAR	31	6.223	0.319	Causality
TAR does not Granger Cause LR		0.249	0.622	No Causality
NB does not Granger Cause TAR	31	0.301	0.588	No Causality
TAR does not Granger Cause NB		9.493	0.005	Causality
DEN does not Granger Cause TAR	31	1.682	0.205	No Causality
TAR does not Granger Cause DEN		5.719	0.024	Causality

Source: Output data from E-views 10.0

In testing the effect of commercial bank regulation on intermediation function, the granger causality analysis was employed. Evidence from Table 13 shows that the liquidity ratio, the number of commercial bank branches, and commercial banks density significantly affect intermediation function defined by private sector credit to RGDP. Causality runs from liquidity ratio, the number of commercial bank branches, and commercial banks density to credit to the private sector ratio to RGDP model of intermediation function at a 5% level of significance. The explanation of intermediation function in terms of total assets of commercial banks to RGDP unveils that liquidity ratio exerts significance on the intermediation function of commercial banks. In furtherance, it was found that intermediation function measured by total assets of commercial banks to RGDP has a significant effect on commercial banks regulation through prime lending rate, the number of commercial bank branches, and commercial banks density. Causality runs from price and geographic regulations to intermediation function at a 5% level of significance.

The determination of the regulation index that most affects the financial intermediation of commercial banks led to variance decomposition estimation. The result in Table 14 discloses that for both models of financial intermediation, it was the number of commercial bank branches that have the most significant influence on the financial intermediation function of commercial banks. This is followed by commercial banks' density and liquidity ratio, while the prime lending rate has the least influence.

Table 14. Variance Decomposition

Financial Intermediation by CPSR						
Period	S.E.	CPSR	PLR	LR	NB	DEN
1	2.111	100.000	0.000	0.000	0.000	0.000
2	2.696	80.261	0.005	0.090	16.805	2.840
3	3.152	70.047	0.022	1.367	26.435	2.130
4	3.497	57.642	0.018	1.266	39.270	1.804
5	3.733	50.616	0.449	1.139	46.191	1.604
6	3.911	46.652	0.907	1.171	49.797	1.473
7	4.029	44.927	1.562	1.438	50.624	1.448
8	4.112	44.366	2.095	1.794	50.214	1.531
9	4.172	44.207	2.488	2.150	49.391	1.763
10	4.218	44.126	2.731	2.429	48.583	2.131
Financial Intermediation by TAR						
1	22.426	100.000	0.000	0.000	0.000	0.000
2	38.753	99.491	0.147	0.007	0.336	0.019
3	49.014	96.776	0.160	1.072	1.005	0.987
4	54.996	91.714	0.128	3.068	2.534	2.557
5	59.008	85.702	0.113	3.846	5.682	4.657
6	62.805	78.700	0.148	3.831	10.664	6.657
7	66.967	71.261	0.211	3.585	16.768	8.175
8	71.536	64.090	0.291	3.282	23.092	9.246
9	76.361	57.704	0.383	2.987	29.030	9.896
10	81.279	52.276	0.483	2.721	34.272	10.248

Source: Output data from E-views 10.0

Table 9 for price regulation expressed by prime lending rate, the first model of intermediation function: private sector credit to RGDP, price, and product regulations have a negative relationship with intermediation function, while a positive relationship was observed for geographic regulation and intermediation function. For the second model of financial intermediation: total assets of commercial banks to RGDP, price and geographic regulation have a negative relationship with intermediation function, whereas product regulation was positively related to intermediation function. The negative relationship between price regulation and intermediation function did not agree with Ezirim and Moughalu (2002), who found a positive relationship between the duo. Though insignificant, it evidences that tightening price regulation through monetary policy adjustments would lower the intermediation function of the commercial banks: a higher interest rate reduces the intermediation function of commercial banks. The granger causality test in Table 13 points out that price regulation has no significant effect on the intermediation function. In other words, the intermediation function of commercial banks is independent of price regulation. This is favorable with Ezirim and Moughalu (2002) that price regulation is not a major policy thrust for monetary regulation in Nigeria, especially when commercial banks are concerned. Conversely, the intermediation function is observed to exert a significant effect on price regulation. In essence, the level of intermediation function of commercial banks determines the price regulation in the economy.

For Table 9 on product regulation: liquidity ratio, product regulation would be said to be positively related with intermediation function with reliance on the second model of financial intermediation, with an adjusted R-squared of 98.50%. This result is consistent with Ezirim and Moughalu (2002) but in total disagreement with Kale et al. (2015) and Zheng et al. (2017) on the negative link between product regulation and intermediation function of banks. Ezirim and Moughalu (2002) assert that the effect of liquidity ratio or reserve requirement may not be necessarily negative because if it affects the supply-side of the intermediation process more than the demand side, the overall effect will have to be positive. The granger causality analysis envisages that product regulation significantly affects intermediation function, thus an important tool of monetary policy influencing intermediation function. Imposing restriction on the activities of commercial banks would in no little magnitude reduce profitability which ultimately affects their intermediation. On the other hand, relaxing restrictions on activities of commercial banks would mean more profitability to the banks by engaging in traditionally non-banking core activities such as insurance, pension, licensing, etc.

Relying on the second model of intermediation function, commercial banks density has a negative relationship with intermediation function, and in tandem with Ezirim and Moughalu (2002), while the number of branches exhibits a positive relationship, though all insignificant. Evidence from Granger causality analysis supports the significant effect of geographic regulation on intermediation function, hence a crucial aspect of commercial banks regulation that should be treated with utmost good faith. When there is an expansion in branch networks of commercial banks, there would be higher financial inclusion as individuals would have access to financial services at affordable prices. The resultant effect would be increased intermediation function owing to the mobilization of funds from large individuals and distributing the same to the deficit units in the economy. The low density of commercial banks results in a lower level of intermediation function as a large fraction of the population are financially excluded: financial intermediation activities would be within the confine of the urban populace.

Conclusion

The intermediation function of the commercial banks: effective and efficient mobilizations and scarce economic resources and allocating same to deficit economic units are pivotal to the growth and development of the economy. In this study, how commercial banks' regulation affects their intermediation function was ascertained. Based on the findings, the study concludes that commercial bank regulation exerts a significant effect on their intermediation function.

In line with the findings, the study makes some recommendations and hopes that it will be giving favorable attention by our policymakers. First, the Central Bank of Nigeria should relax price regulation through a reduction in the monetary policy rate. Relaxation of price regulation will increase credit requests from productive segments of the economy, thus a rise in intermediation activities of commercial banks. Secondly, product regulation should be relaxed too to allow commercial banks to engage in various conventionally non-banking activities to cause an upsurge in the level of their intermediation, and the resultant effect on the part of commercial banks be a rise in profitability. Finally, apart from capital adequacy and reserve requirements, other pre-requisite such as documentation requirements for establishing new branches and other banking outlets should be relaxed to allow banks to open more branches and enter the market. This will, to a very great extent, boost the intermediation function of commercial banks.

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Government fiscal spending and crowd-out of private investment: An empirical evidence for India

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Abstract

Purpose — The paper evaluates the crowding-in or crowding-out relationship between public and private investment in India, controlling fiscal and monetary variables.

Methods — In a flexible accelerator theoretical framework, the paper estimates long and short-run investment dynamics, employing Autoregressive Distributed Lag (ARDL) cointegration approach. We use a back series of national account statistics that incorporates enhanced coverage of the organized corporate sector.

Findings — Our results suggest investment complementarity between the public and private sector at an aggregate and sectoral level over the period 1981-2019. Barring short-run crowding-out in construction and financial services at industry level, public investment stimulates private counterparts, both in the long and short-run. However, fiscal deficit, inflation expectation, and sovereign vulnerability influence private investment adversely. Moreover, the long-run crowding-out bearing of fiscal imbalance is quantitatively higher when the public sector invests in mining and manufacturing and insignificant with infrastructure.

Implication — Sizable infrastructure investment as a proportion of government finances would moderate the adverse impact of the deficit on private investment. Further, quality fiscal adjustments and containing inflation would enhance private investment activities.

Originality — Besides aggregate and sectoral levels, the study also evaluates the impact of industry-level public investment on private capital expenditure. This paper also incorporates derived variables in the regression framework using statistical filters and the principal component technique.

Keywords — Public investment, private investment, crowding-out, statistical filter.

Introduction

Economic growth of an economy is contingent upon investments in physical capital, human capital and technological development by all its stakeholders. In this respect, crowding-out of private investment attracts attention time and again, more so with an expansionary fiscal action in developing countries. Broadly, crowding-out is expected when public investment utilises more resources that would be otherwise available to the private sector or if it produces marketable output that competes with the output produced by the private sector. The 'real or direct' crowd-out occurs

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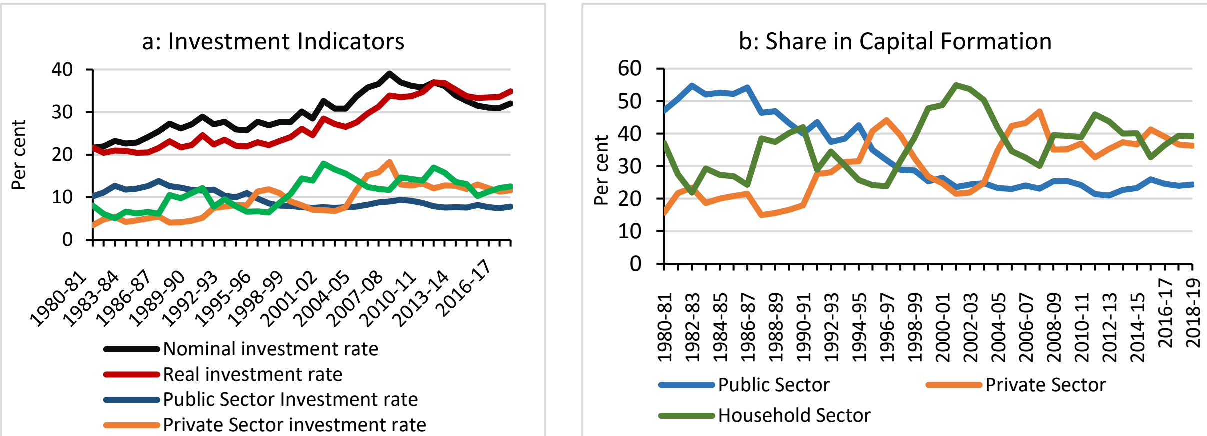
when increase in public capital formation displaces private investment directly; whereas partial loss of private sector investment due to hardening of interest rates is referred as 'financial or indirect' crowd-out (Buiter, 1990). For developing economies, enhanced public investment also supports growth indirectly through positive sentiment building channel and helps productivity enhancement leading to healthier growth prospects (Afonso & St. Aubyn, 2010). Though the country specific factors dominate while making investment decision by private corporates, public fixed capital formation in developing economies has virtuous complementary effect (Atukeren, 2010). On a similar line, considering 39 developed and developing economies, Ahmed and Miller (2000) conclude that government spending in transport and communication attracts private investment only in developing countries, while expenditure on social security and welfare scheme crowd-out investment for both developed and developing economies.

In country specific studies, government spending crowd-in private investment in Turkey, while budget deficits crowd it out (Şen, H., & Kaya, 2014). Moreover, Wang (2005) argues that public expenditure on health and education in Canada compliments, while expenditure on infrastructure and social security schemes substitutes private investment due to active participation of private corporates in physical infrastructure sector than in human capital formation. The asymmetric substitutive and complimentary relationship have also been observed for the Brazilian and Chinese economy in short and long-run, respectively (Cruz & Teixeira, 1999; Wu & Zhang, 2009). However, for South Africa, government investment neither crowds-in nor crowds-out private investment, though government spending in social and infrastructure sectors may indirectly influence private investment through accelerator effect (Kollamparambil & Nicolaou, 2011). In case of Pakistan, studies opinion that government's expenditure in defence, debt servicing and manufacturing sector decelerates private investment, while expenditure on infrastructure, health, education and agriculture attracts (Hussain, Muhammad, Akram, & Lal, 2009; Saeed, Hyder, Ali, & Ahmad, 2006). In Indian context, government's investment driven short-run substitutive impact for private sector is neutralised through output expectations in long run, thereby offsetting initial crowding-out effect (Dash, 2016; Mitra, 2006). Serven (1999) however, at sectoral level concludes that public infrastructure investment complements private investment in long-run, but crowds-out in short-run through credit rationing and hardening of interest rate. Contrary to this, using asymmetric VAR, Chakraborty (2007) opines investment complementarity conditional upon exchange rate and fiscal deficit between the two sectors. Moreover, underscoring importance of reforms, Bahal, Raissi, and Tulin (2018) argue for complimentary relationship only with restricted dataset, post 1980. Recently, estimation using non-linear ARDL also argues investment complementarity for India in long and short-run controlling foreign direct investment, expected output and interest rate (Akber, Gupta, & Paltasingh, 2020).

The fiscal deficit influences adversely to private investment and poses a downside risk to growth. Many studies considering panel data suggest that deficit crowds out private investment marginally in developed countries (Mahmoudzadeh, Sadeghi, & Sadeghi, 2017) and significantly in oil-dependent economies primarily through credit channel than interest rate channel (Anyanwu, Gan, & Hu, 2018). In a similar line, fiscal deficit significantly crowds-out private investment for South Africa (Biza, Kapingura, & Tsegaye, 2015), India (Chakraborty, 2016; Dash, 2016), China (Wu & Zhang, 2009), and Pakistan (Hussain et al., 2009). Also, trade openness and output cyclicality have positive on investment for developing economies (Furceri & Sousa, 2011).

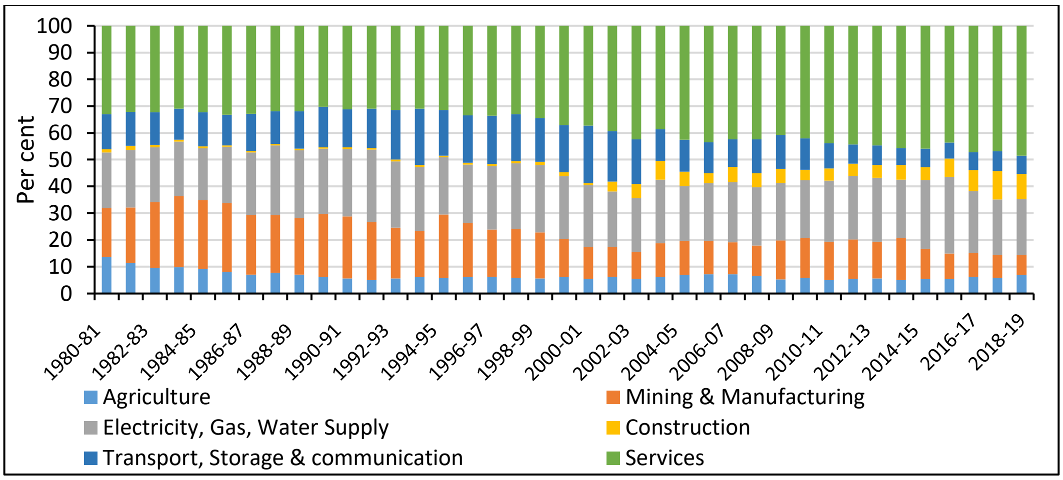
The literature could not however support a single view favouring either complementary or substitutive effect of public investment on private capital formation. Moreover, theoretical foundation, empirical framework, domestic factors and granularity analysis point to dissimilar conclusion. Since, the crowd-out hypothesis of private investment is still inexplicable for Indian economy, we revisit and display annotations of dynamic link between the gross fixed capital formation (GFCF) of public and private sectors with top-down approach – starting with aggregate level and drill down to industry level – controlling macro-economic, monetary and fiscal variables. Primarily the study analyses direct crowd-out effect, interest rate dynamics is captured through credit cost, inflation expectation besides macroeconomic vulnerability and economic cycle. Accordingly, the explicit hypothesis of this study are (a) public investment crowds-out private investment at aggregate, sectoral (infrastructure, non-infrastructure, services etc.) and industry level, (b) fiscal deficit crowds-out private investment, and (c) macroeconomic vulnerability and expected inflation dampen private investment. In this sense, the study aims to contribute to the literature by analysing empirically the ambiguous crowding-out hypothesis for India considering comprehensive macro-economic and fiscal variables at aggregate, sectoral and industry level. Such comprehensive consideration of probable investment's determinants are missing in literature for country specific study.

The GFCF is acquisition and creation of assets for their use in medium to long run netted against disposals of produced fixed assets. In Indian context, this study makes probably first use of back series disaggregated (industry level) investment statistics covering financial year 1980-1981 to 2018-2019 with 2011-12 as reference year. The public sector investment rate[1] was higher than the private sector until early 1990s. But the trend reversed by mid-1990s; private investment crossed over and remains above public investment barring few intermittent years in early 2000s (Figure 1a). However, moderation of investment since 2006-07 is a concern, though the recent broad basing of credit uptick may streamline sectoral investments going forward. Declining of stalled projects, reforms in telecom, power, real estate and synchronised public-private partnerships for infrastructure help investment revival (Raj, Sahoo, & Shankar, 2018).



Source: Reserve Bank of India

Figure 1. Investment Dynamics



Source: Government of India

Figure 2. Share of Public Sector Investment

[1] Public investment includes government investment also. Investment rates are the ratio of GFCF to the gross value addition (GVA).

The investment share of private sector crossed above public sector and remained elevated since 2003-04 (Figure 1b), however, gap between public investment in infrastructure and non-infrastructure has been widening in recent years (Figure 2). The decadal correlation coefficient between public and private investment rate revels that initial substitutive effect (correlation -0.61 during 1990-2000) reverses to complementarity (correlation 0.78 and 0.70, respectively) during subsequent decades (Table 1). However, for entire study period negative correlation (-0.65) is suggestive of a substitutive relationship. Additionally, volatility of public investment rate (1.9) was about half that of private investment (3.8), points toward unreliable nature of private investment in Indian context.

Table 1. Investment Rate - Stylised Summary

	Private Sector				Public Sector				
Years$	Mean	Max	Min	Std. Dev.	Mean	Max	Min	Std. Dev.	Correlation
1980-81 to 1989-90	4.6	5.5	3.4	0.7	12.1	13.8	10.2	1.0	0.72**
1990-91 to 1999-00	8.8	11.9	5.2	2.0	9.7	11.8	7.7	1.5	-0.61*
2000-01 to 2009-10	11.5	18.3	6.8	4.2	8.3	9.4	7.5	0.8	0.78***
2010-11 to 2016-17	12.3	13.2	11.3	0.6	7.8	8.7	7.4	0.4	0.70**
1980-81 to 2018-19	9.2	18.3	3.4	3.8	9.5	13.8	1.9	1.9	-0.65***

Notes: ***, ** and * represents significance at 1%, 5 % and 10% level, respectively.
$: Financial year of India (April to March)

Whether deficit financing poses capital constraint to private investment is a much-debated issue of macro-economy; more so in developing economy like India considering recent fiscal slippage and dismal investment scenario amidst subsidy support, bank's asset quality deterioration, squeeze of credit flows to non-financial private sector and protracted slowdown. Therefore, examining crowding-out phenomenon is essential for informed policy decision to mitigate potential adverse impact on productive private investment.

Methods

Following Jorgenson (1967) and Blejer and Khan (1984), the desired capital stock may proportional to the expected output.

$$KPvt_t^* = \propto Y_t^{exp} \tag{1}$$

where, $KPvt_t^*, Y_t^{exp}$ and $\propto$ are desired capital stock by private sector, expected output and a constant, respectively. Further, under partial adjustment mechanism, the actual stock of capital may be adjusted with an adjustment factor to the difference between desired level of capital and actual stock at previous year.

$$\Delta KPvt_t = \beta(KPvt_t^* - KPvt_{t-1}) \tag{2}$$

where, $KPvt$ is the actual capital stock, $\Delta KPvt$ net private investment and β (0≤ β ≤ 1) adjustment coefficient. Theoretically gross private investment is sum of the net private investment and depreciation of the previous capital stock i.e.

$$IPvt_t = \Delta KPvt_t + \delta KPvt_{t-1} \tag{3}$$

$$IPvt_t = \beta(KPvt_t^* - KPvt_{t-1}) + \delta KPvt_{t-1} \tag{4}$$

$$IPvt_t = [1-(1-\delta)L]KPvt_t\ ;\ LKPvt_t = KPvt_{t-1} \tag{5}$$

Where, $IPvt_t$ is gross private investment and δ is depreciation rate. After little algebraic rearrangement, we get following representation for $IPvt_t$.

$$IPvt_t = [1-(1-\delta)L]\beta\ KPvt_t^* + (1-\beta)\ KPvt_{t-1} \tag{6}$$

By basic assumption of the flexible accelerator model as per equation (1), the private investment may be expressed as:

$$IPvt_t = \beta \propto [1 - (1-\delta)L]\, Y_t^{exp} + (1-\beta)\, KPvt_{t-1} \tag{7}$$

The coefficient β captures the private investment response to the gap between desired and actual investment. It varies with factors influencing investment ability of private sector to achieve the desired level of investment.

$$\beta = \psi\ (IPub, GFD, WALR, Inf, credit, OG, trade, vulnerability)$$

In general, the investment activities depend on credit availability, lending rate and output gap[2] (Peltonen, Sousa, & Vansteenkiste, 2012); and expected inflation particularly for inflation targeting economies (Grasso & Ropele, 2018). The importance of uncertainty in investment as argued by Güney (2020) has been captured through sovereign macroeconomic vulnerability (SMV) - a 3-year rolling standard deviation of cyclic components of vulnerability indicator[3]. Besides, to capture the capital flow investment dynamics, we consider trade openness. The dynamic linear representation of private investment is therefore framed as below.

$$PRC_t = f(PRC_{t-i}, PS_t, GFD_t, WALR_t, Credit_t, OG_t, TO_t, SMV_t, Inf_t);\ i = 1,2, \dots n$$

We employ autoregressive distributed lag (ARDL) approach (Pesaran, Sin, & Smith, 2001) for its implicit advantages over traditional co-integrating framework as it takes care endogeneity of explanatory variables. Both the short and long-run relationship is estimated simultaneously in presence of I(1) variables through the bound test. Further, the Error Correction Model (ECM), derived through a linear transformation, integrates short-run adjustments with long-run equilibrium without losing long-run information.

The investment and macroeconomic statistics are sourced from National Statistical Organization (NSO), Government of India (GoI) and Handbook of Statistics on Indian Economy (HBS), Reserve Bank of India (RBI), respectively. The derived variables *viz.*, output gap, sovereign macroeconomic vulnerability and expected inflation are estimated using statistical filters and principal component technique (Appendix). The variables description are given in Annexure Table 1. A priori, we expect a positive or negative coefficient of public investment and output gap in determining private investment. The banks credit flow and trade openness are expected to interact positively; however, lending rate, expected inflation, fiscal deficit and sovereign vulnerability may have a negative coefficient.

Results and Discussion

We carry out ARDL bound test to examine long-run cointegrating relationship using unrestricted ECM controlling variables stated above. We first check for stationarity of these variables using Augmented Dickey-Fuller (ADF) and Philips-Peron (PP) tests with constant and time trend. Witnessing many reforms in India during the study period, we also report breakpoint ADF test following (Perron, 1989)[4]. The output gap (OG), non-food credit flow (NCF), lending rate (WALR), trade openness (TO) and in few instances public sector capital formation at industry level are stationary only with structural break (Annexure Table 2). We may thus conclude that all variables are either I(0) or I(1) with or without breakpoint. The comprehensive representation of ARDL long-run cointegration model, considering all probable variables impacting private investment, is specified below before testing the joint significance of lagged variables using F-test (Wald test).

$$\begin{aligned}\Delta \ln PRC_t = \mu &+ \textstyle\sum_{i=1}^{n} \beta_{1i}\, \Delta \ln PRC_{t-i} + \sum_{i=1}^{n} \beta_{2i}\, \Delta \ln PS_{t-i} + \\ &\textstyle\sum_{i=1}^{n} \beta_{3i}\, \Delta ln\ GFD_{t-i} + \sum_{i=1}^{n} \beta_{4i}\, \Delta OG_{t-i} + \sum_{i=1}^{n} \beta_{5i}\, \Delta ln\ NCF_{t-i} + \sum_{i=1}^{n} \beta_{6i}\, \Delta WALR_{t-i} + \\ &\textstyle\sum_{i=1}^{n} \beta_{7i}\, \Delta ExpINF_{t-i} + \sum_{i=1}^{n} \beta_{8i}\, \Delta TO_{t-i} + \sum_{i=1}^{n} \beta_{9i}\, \Delta SMV_{t-i} + \varphi_1 \ln PRC_{t-1} +\end{aligned}$$

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[2] Output Gap (OG) = (Actual Output-Potential Output)/ Potential Output *100

[3] Vulnerability Indicator = (Fiscal Deficit/GDP+ Current Account Deficit/GDP+ GDP deflator)

[4] We consider models for data having trend with one-time break determined endogenously assuming break occurs gradually and follows same dynamic path as the innovations (innovation outlier). The tests evaluate the null hypothesis of a unit root process, possibly with a break, against alternative hypothesis of trend stationary with break.

$$\varphi_2 \ln PS_{t-1} + \varphi_3\, GFD_{t-1} + \varphi_4\, OG_{t-1} + \varphi_5 \ln NCF_{t-1} + \varphi_6\, WALR_{t-1} + \varphi_7\, ExpINF_{t-1} + \varphi_8\, TO_{t-1} + \varphi_9\, SMV_{t-1} + \rho Dummy_i + \epsilon_t \quad (8)$$

To measure the asymmetric impact on private investment at sectoral level of public investment (PS_t), we segregate public investment in infrastructure, non-infrastructure, core infrastructure and service sector, considering importance of infrastructure investment as it leads to employment driven growth for emerging economies (Nasution & Imam, 2017). We consider public non-infrastructure investment as investment in mining & quarrying, manufacturing, trade, hotels & restaurants, financial services, real estate and business services, public administration, defence and other services (Chakraborty, 2016). Similarly, infrastructure investment includes investment in agriculture (especially irrigation); electricity, gas, water supply; construction; transport, storage and communication. Further, the core-infrastructure excludes agriculture from infrastructure investment as capital formation under irrigation projects is not explicit. We further extend the analysis at industry level of public sector investment. The null hypothesis of no cointegrating relationship is as below:

$$H_0: \varphi_j = 0\ for\ j = 1,2, \dots\dots, n$$

The F-statistic (Wald test) under the null follows a non-standard distribution in presence of non-stationary variables and thus test statistic is compared with two asymptotic critical values in presence of I(0) and I(1) variables to conclude a cointegrating relationship (Pesaran et al., 2001). If computed F-statistic falls outside critical value of I(1), we may reject the null; if falls within the critical band of I(1) and I(0), decision is inconclusive; and outside with reference to I(0) critical value implies no cointegrating relationship among the covariates. Subsequently, short run ECM representation is estimated as below:

$$\Delta \ln PRC = \alpha + \sum_{i=1}^{n} \beta_{1i}\, \Delta \ln PRC_{t-i} + \sum_{i=1}^{n} \beta_{2i}\, \Delta \ln PS_{t-i} + \sum_{i=1}^{n} \beta_{3i}\, \Delta ln\, GFD_{t-i} + \sum_{i=1}^{n} \beta_{4i}\, \Delta OG_{t-i} + \sum_{i=1}^{n} \beta_{5i}\, \Delta ln\, NCF_{t-i} + \sum_{i=1}^{n} \beta_{6i}\, \Delta WALR_{t-i} + \sum_{i=1}^{n} \beta_{7i}\, \Delta ExpINF_{t-i} + \sum_{i=1}^{n} \beta_{8i}\, \Delta TO_{t-i} + \sum_{i=1}^{n} \beta_{9i}\, \Delta SMV_{t-i} + ECM_{t-1} + \rho Dummy_i + \vartheta_t \quad (9)$$

Where, Δ denotes difference operator, β indicates short run multiplier of respective regressors and ECM_{t-1} is lagged error correction term with expected coefficient ranged between (-1, 0).

Table 2. ARDL Bounds Test at Aggregate Level of Investment

ARDL Models (Lag)	2,0,0,1,1,2,0,0	2,1,0,1,1,2,0,0	2,2,0,1,1,2,0,0	2,0,0,1,2,2,0,0
Dependent Variable: Private Corporate Sector Investment (Ln PRC)				
Common Regressors	Ln PS, Ln GFD, Ln NCF, WALR, EINF			
Distinct Regressors	OG, TO	TO, SMV	OG, SMV	SMV, DREF*LPS
F-statistic	10.87***	10.85***	12.89***	16.34***
Diagnostic Statistics				
$\boldsymbol{F-Stat_{Serial[2]}}$	0.7651 (0.27)	3.8873 (0.11)	4.2554 (0.13)	2.6207 (0.10)
$\boldsymbol{F-Stat_{Arch[1]}}$	0.0090 (0.92)	1.3395 (0.65)	1.9992 (0.77)	1.2732 (0.29)
$\boldsymbol{F-Stat_{Reset[1]}}$	1.1671 (0.29)	0.5397 (0.47)	0.0969 (0.76)	0.4402 (0.51)
Critical Value Bounds for Regressors (k=7)				
Significance Level	10%	5%	2.50%	1%
Lower Bounds I(0)	1.92	2.17	2.43	2.73
Upper Bounds I(1)	2.89	3.21	3.51	3.90

Notes: ***, ** and * denote statistical significance at 1%, 5 % and 10 % level, respectively. p-value are reported in parenthesis. Ln denotes natural logarithm. Square brackets represent lag length.

As the computed F- statistic falls in the rejection region, the bound test confirms a cointegrating relationship among covariates (Table 2). The results at aggregate level are presented in Table 3. The long run estimated coefficients of public investment is positive (ranged from 0.78 to 1.14) and statistically significant across the models; indicate a complementarity between private and public investment; in line with the previous studies of Muthu (2017), Bahal et al. (2018) and (Akber et al., 2020) for India. The net credit flow and cost of credit impact positively, imply that private investment in long-run is insensitive to interest rate with adequate credit flow. However, expected inflation dampens the private investment. Moreover, deficit coefficient signifies that one per cent increase of fiscal deficit reduces private investment by over half percent. The estimated trade openness coefficient (-0.03) have diminishing effect on the private investment, possibly due to persistency of trade deficit over the years for India economy.

The negative coefficients of inflation expectation and sovereign vulnerability underscore their importance for private sector investment in long-run. Thus, for policy perspective, containing inflation and vulnerability are imperative to buttress private investment. The negative elasticity of output gap, though low indicates the slowdown of private investment when economy is overheated (positive output gap). Thus, optimizing the actual output around the potential may be desirable to attract private investment. Next, we estimate the short-run dynamics with appropriate exogenous dummies for 1990s reform, structural breaks and general election year of India for capturing political uncertainty in election year. The estimated ECM coefficients are negative, statistically significant and bounded by (-1, 0) (Table 4). It signify that against the short-term disequilibrium, private investment adjusts in a range of 81 to 89 per cent within a year to restore long-run equilibrium. In short-run, cost of credit and expected inflation dampen private investment, but credit flow has significant positive relationship as expected. The reform dummy is significant and positive, but election dummy doesn't diminish private investment against the conjecture.

Table 3. Long Run Estimates at Aggregate Level

Dependent Variable: Private Corporate Investment (Ln PRC)				
ARDL Model/Variables	2,0,0,1,1,2,0,0	2,1,0,1,1,2,0,0	2,2,0,1,1,2,0,0	2,0,0,1,2,2,0,0
Ln PS	0.779***(5.799)	0.906***(11.557)	1.136***(10.013)	1.054***(8.812)
Ln GFD	-0.538*** (-3.741)	-0.518***(-4.95)	-0.556***(-4.32)	-0.562***(-4.522)
Ln NCF	0.71***(6.633)	0.554***(7.846)	0.373***(8.217)	0.476***(7.055)
WALR	0.057*(2.048)	0.05***(3.139)	0.056**(2.596)	0.084***(3.252)
EINF	-0.467***(-5.446)	-0.358***(-5.365)	-0.308***(-4.658)	-0.306***(-4.073)
OG	-0.042*(-1.861)		0.0070 (0.504)	
TO	-0.029***(-3.029)	-0.015**(-2.656)		
SMV		-0.069***(-3.019)	-0.096***(-3.745)	-0.070***(-3.314)
DREF*Ln PS				0.020**(2.732)
C	4.603** (2.387)	3.351**(2.499)	2.045 (1.636)	1.771 (1.046)

Notes*:* PRC denotes private corporate sector investment. The figures in parentheses are t-statistics.
***, ** and * denote statistical significance at 1%, 5 % and 10 % level, respectively. Ln denotes natural logarithm.

Table 4. Short Run Estimates at Aggregate Level

Dependent Variable: Δ Private Corporate Investment (Δ Ln PRC)				
Δ Ln PRC (-1)	0.245***(3.581)	0.245***(3.671)	0.181***(2.92)	0.189***(3.384)
Δ Ln PS		1.162***(7.936)	1.586***(10.142)	
Δ Ln PS(-1)			0.416**(2.486)	
Δ Ln NCF	0.253***(7.236)	0.176***(5.391)	0.078***(3.127)	0.146***(5.106)
Δ WALR	-0.097***(-3.845)	-0.104***(-4.102)	-0.079***(-3.453)	-0.103***(-4.791)
Δ WALR(-1)				-0.050**(-2.389)
Δ EINF	-4.088***(-10.161)	-3.421***(-10.089)	-2.574***(-10.336)	-2.658***(-11.379)
Δ EINF(-1)	4.037***(10.051)	3.402***(10.237)	2.531***(10.475)	2.628***(11.224)
DBREAK	0.072*(1.853)		-0.057*(-1.888)	
DREF		0.250***(8.868)	0.386***(11.639)	
DELECTION			0.097***(4.349)	
ECM t-1	-0.814***(-11.549)	-0.872***(-11.613)	-0.864***(-12.944)	-0.894***(-14.162)
		Diagnostic Statistics		
Adjusted R^2	0.76	0.83	0.88	0.83
S.E. of regression	0.082	0.069	0.059	0.070
SBC	-1.68	-1.96	-2.098	-2.000
DW Statistics	2.23	2.74	2.77	2.56
F-Statistic	864.31(000)	1124.4(000)	1235.51(000)	1190.62(000)

Note: 'Δ' indicates differenced operator. The figure in parenthesis denotes t-statistics. ***, ** and * denote statistical significance at 1%, 5 % and 10 % level, respectively. In F-Statistics, the figure in parenthesis indicates the p-value. Ln denotes natural logarithm.

Table 5. Long Run Estimates with Public Sector Investment at Sectoral Level

Dependent Variable: Private Corporate Sector Investment (Ln PRC)							
ARDL Model	**2, 0, 0, 1, 2, 2, 0**	**2, 0, 0, 1, 2, 2, 0, 2**	**2, 0, 0, 1, 2, 2, 0**	**2, 0, 0, 1, 2, 2, 0, 2**	**2, 0, 0, 1, 2, 2, 0**	**2, 0, 0, 1, 2, 2, 0, 0**	**2, 1, 0, 1, 1, 2, 0, 0**
Ln GFD	-0.538***(-5.893)	-0.51***(-4.844)	-0.665***(-11.26)	-0.655***(-5.94)	-0.576***(-14.591)	-0.55***(-7.804)	-0.564***(-6.031)
Ln NCF	0.604***(12.245)	0.563***(8.036)	0.499***(5.864)	0.385***(3.62)	0.647***(14.605)	0.649***(14.315)	0.545***(7.095)
WALR	0.084***(4.717)	0.073***(3.015)	0.129***(3.86)	0.100***(3.531)	0.074***(4.396)	0.075***(3.967)	0.092***(3.809)
EINF	-0.299***(-4.407)	-0.325***(-3.501)	-0.484***(-4.819)	-0.57***(-5.13)	-0.289***(-4.45)	-0.28***(-3.606)	-0.297***(-3.32)
SMV	-0.083***(-4.563)	-0.090***(-3.624)	-0.086***(-4.025)	-0.088***(-3.622)	-0.083***(-4.302)	-0.087***(-3.895)	-0.074***(-3.001)
OG		-0.0180 (-0.569)		-0.058*(-1.859)		0.0110 (0.823)	
TO							-0.019***(-3.697)
Ln INFRA	0.930***(10.421)	0.916***(10.442)					
Ln Non-INFRA			1.082***(14.325)	1.116***(9.137)			
Ln Core -INFRA					0.891***(15.182)	0.868***(14.314)	
Ln Service							0.977***(16.04)
C	2.414 (1.646)	3.065 (1.596)	4.102*(1.726)	5.944**(2.662)	3.024**(2.144)	2.867*(1.723)	3.134 (1.579)
F-statistic (Bound test)	15.6041***	13.4508***	14.1650***	13.8450***	13.8787***	12.0092***	10.7602 ***

Note: ***, ** & * indicate 1%, 5 % and 10% level of significance, respectively. The figures in parentheses are t-statistics. Ln denotes natural logarithm.

Table 6. Short Run Estimates with Public Sector Investment at Sectoral Level

Dependent Variable: Δ Private Corporate Sector Investment (Δ Ln PRC)							
Δ Ln PRC(-1)	0.256***(4.328)	0.278***(4.805)	0.162**(2.496)	0.185***(3.087)	0.254***(4.079)	0.275***(4.512)	0.161**(2.683)
Δ Ln NCF	0.236***(7.059)	0.223***(6.868)	0.118***(3.67)	0.074**(2.484)	0.248***(6.977)	0.251***(7.089)	0.131***(4.179)
Δ WALR	-0.121***(-5.037)	-0.14***(-5.681)	-0.08***(-3.359)	-0.101***(-4.391)	-0.128***(-4.982)	-0.133***(-5.191)	-0.065***(-2.899)
Δ WALR(-1)	-0.071***(-3.024)	-0.072***(-3.111)	-0.065**(-2.681)	-0.054**(-2.409)	-0.068**(-2.747)	-0.072***(-2.928)	
Δ EINF	-2.943***(-10.6)	-3.274***(-11.139)	-3.001***(-10.169)	-3.759***(-11.631)	-2.84***(-9.921)	-2.823***(-9.989)	-2.63***(-9.652)
Δ EINF(-1)	2.879***(10.441)	3.241***(11.066)	3.016***(10.095)	3.808***(11.624)	2.738***(9.726)	2.735***(9.804)	2.69***(9.941)
Δ OG		0.0010 (0.074)		-0.0030(-0.259)			
Δ OG(-1)		0.029**(2.517)		0.041***(3.556)			
Δ Ln Service							1.221***(9.423)
Δ DREF							0.241***(8.453)
ECM t-1	-0.864***(-12.76)	-0.89***(-13.018)	-0.721***(-12.158)	-0.771***(-13.208)	-0.84***(-12.034)	-0.844***(-12.14)	-0.756***(-11.564)
Diagnostic Statistics							
Adjusted R^2	0.83	0.80	0.77	0.81	0.77	0.77	0.85
S.E. of regression	0.07	0.075	0.08	0.074	0.081	0.08	0.07
SBC	-2.304	-1.743	-1.740	-1.767	-1.723	-1.738	-2.106
DW Statistics	2.56	2.60	2.50	2.70	2.51	2.58	2.57
$F - Stat_{Serial}$ [2]	2.5584 (0.11)	2.4777 (0.11)	3.3697 (0.05)	6.8545(0.06)	2.0774 (0.15)	2.4327 (0.15)	2.5317 (0.11)

We further analyze the impact of public investment at sectoral level for private investment activities. The ARDL bound tests confirm the existence of a cointegrating relationship among the covariates. The long and short-run coefficients are presented in Table 5 and Table 6, respectively. The results confirm that (a) fiscal deficit dampens private investment in long-run and more severe when public sector invests in non-infrastructure/service sector, (b) uptick in credit flow promotes private investment in long run, signifies a conducive credit environment for spurring the investment growth, (c) inflation expectation bearings negatively for private investment in long and short-run though lending rate pulls it down in short-run only, (d) sovereign vulnerability squeezes private investment in long run across the sector. Further, public sector investment in infrastructure, non-infrastructure, core-infrastructure and services sector compliment homogeneously to private capital formation. However, trade openness doesn't help to boost private investment in long-run (coefficient: -0.19) indicative of reliance on domestic investment, though 1990s reform helps to spur private investment in short-run (coefficient: 0.24) in India.

The crowding-in results of infra and non-infra investment at sectoral level are consistent with empirical evidence of Chakraborty (2016) for India and Kollamparambil and Nicolaou (2011) for South Africa. The result however, contradicts insignificant impact of Muthu (2017) for infra and non-infra sector in India. Further, none of the previous study empirically adjudged crowd-out effect of public investment in service industries against perception of a substitutive effect. This study observes a positive elasticity of public service sector investment and hence supports for private investment. Moreover, the negative impact of fiscal deficit is quantitatively higher when public sector invest in non-infra or service sector as higher government borrowing may push either interest rate up or displace investable resources away from private corporate.

We also analyze whether public sector investment at industry level impacts private investment differently; as literature is scanty at this granularity. The ARDL bound tests confirm the existence of cointegrating relationship. The estimated long-run coefficients of industry wide public investment (0.28 to 1.02) signifies complementarity with private sector (Table 7), consistent with aggregate and sectoral crowding-in results. Broadly, results are consistent with Muthu (2017) for most of the core-infrastructure industries for India as against crowd-in effect observed across the industries in this study. Further, the long-run negative bearing due to deficit is more significant with public investment in mining and manufacturing (coefficient -1.14) and insignificant with construction and transport, storage and communication (infrastructure sector). Such negative deficit-investment nexus was observed for most of the emerging economies (Biza et al., 2015; Chakraborty, 2016; Wu & Zhang, 2009). Credit flow supports investment, but inflation and macroeconomic vulnerability drag it in the long and short-run. The cost of credit also dampens investment prospects in short-run, in-sync with findings at aggregate and sectoral level. The estimated ECM coefficients indicate reversion to long-run stability consistently; adjust 35 to 77 per cent within a year's time (Table 8). However, in short-run, barring public investment in financial services and construction, other industry level investment supports private capital formation. Active investment of corporate sectors in financial and construction industries; and resource constraints probably explain the short-run substitutive effect (instantly and with a year lag, respectively) in these industries.

Robustness checks of the estimated models have been carried out for serial correlation, error homoscedastic structure and modal specification; and reported appropriately. Moreover, CUSUM and CUSUM square plots also confirm the parametric stability at 5 per cent level of significance.

Table 7. Long Run Estimates with Public Sector Investment at Industry Level

Dependent Variable: Private Corporate Sector Investment (Ln PRC)						
ARDL Model	**1,1,1, 0,1,1,2,2**	**2, 1, 0, 1, 2, 2, 0**	**1, 2, 1, 1, 2, 0, 0**	**1,1,2,1,1,2,0,0**	**2,2,1,2,1,2,2,2**	**1, 1, 2, 1, 2, 0, 0**
Ln GFD	-1.137***(-7.249)	-0.638***(-6.31)	-0.1920 (-1.079)	-0.1590 (-1.029)	-0.498***(-3.682)	-0.0980 (-0.902)
Ln NCF	0.449***(4.983)	0.688***(6.247)	0.823***(5.469)	0.702***(5.681)	0.646***(4.993)	1.038***(12.002)
WALR	0.124**(2.635)	0.092**(2.334)	0.396***(6.205)	0.326***(6.044)	0.113**(2.542)	0.243***(9.744)
EINF	-0.865***(-6.271)	-0.382***(-3.573)	0.111*(1.834)	0.149**(2.617)	-0.1190 (-0.886)	0.159***(3.115)
SMV	0.0670 (1.44)	-0.082**(-2.76)	-0.094*(-1.799)	-0.191***(-3.131)	0.0510 (1.372)	-0.086***(-2.952)
Ln MINI	1.019***(4.494)					
Ln MANU	0.757***(3.208)					
Ln EGW		0.824***(7.529)				
Ln CON			0.567**(2.653)	0.516***(3.961)		
Ln TSC				0.278*(1.922)		
Ln FS					0.444**(2.326)	
Ln THR					0.849***(8.138)	
Ln RES						0.355***(5.898)
C	8.792**(2.353)	5.043**(2.106)	-5.199***(-3.942)	-5.971***(-5.321)	1.9720(0.73)	-3.977***(-5.244)
F-statistic (Bound test)	12.515***	8.505***	13.586***	12.370***	8.491***	12.991***
Diagnostic Statistics						
$F-Stat_{Serial}$ [2]	1.519 (0.25)	1.051 (0.36)	1.481 (0.25)	3.278 (0.11)	0.477206 (0.63)	0.820 (0.45)
$F-Stat_{Arch}$[1]	0.0262 (0.87)	0.006872 (0.93)	0.002 (0.97)	0.100 (0.75)	0.733165 (0.40)	0.560 (0.46)
$F-Stat_{Reset}$[1]	0.336 (0.57)	0.026243 (0.87)	0.023 (0.88)	0.223 (0.64)	0.248089 (0.81)	2.068 (0.16)

Note: The figures in parentheses are t-statistics. ***, ** & * indicate 1%, 5 % and 10% level of significance, respectively. Ln denotes natural logarithm. In F-Statistics, the figure in parenthesis indicates the p-value. Square brackets represent lag length.

Table 8. Short Run Estimates with Public Sector Investment at Industry Level

Dependent Variable: Δ Private Corporate Investment (Δ Ln PRC)						
Δ Ln PRC(-1)		0.196**(2.478)			-0.212**(-2.35)	
Δ Ln NCF	0.097**(2.607)	0.234***(5.845)	0.128***(3.286)	0.133***(4.194)	0.208***(7.143)	0.23***(5.967)
Δ WALR	-0.086***(-3.144)	-0.128***(-4.221)	0.0130(0.486)	0.0170(0.771)	-0.081***(-3.538)	-0.108***(-3.567)
Δ WALR (-1)		-0.079**(-2.796)	-0.064**(-2.204)	-0.077***(-3.189)	-0.042*(-2.058)	-0.116***(-3.84)
Δ EINF	-2.359***(-9.743)	-2.868***(-8.15)			-1.814***(-7.587)	
Δ EINF(-1)	2.302***(9.332)	2.826***(8.088)			2.344***(8.38)	
Δ SMV	-0.062***(-3.131)				-0.036**(-2.584)	
Δ SMV(-1)	-0.053**(-2.725)				-0.05**(-2.84)	
Δ Ln GFD			-0.393***(-4.753)	-0.359***(-5.412)	-0.543***(-9.486)	-0.403***(-4.944)
Δ Ln GFD (-1)					-0.262***(-3.157)	-0.231**(-2.513)
Δ Ln Mining	0.366***(5.471)					
Δ Ln Manu	0.0450 (0.5)					
Δ Ln EGW		0.347***(2.941)				
Δ Ln CON			0.0330 (0.793)	0.0380 (1.137)		
Δ Ln CON (-1)			-0.134***(-2.955)	-0.114***(-3.165)		
Δ Ln TSC				0.500***(5.826)		
Δ Ln FS					-0.0750 (-1.683)	
Δ Ln FS (-1)					-0.222***(-3.655)	
Δ Ln THR					0.339***(7.898)	
Δ Ln RES						0.0570 (1.062)
ECM t-1	-0.575***(-12.558)	-0.766***(-9.47)	-0.353***(-11.907)	-0.42***(-12.399)	-0.748***(-10.825)	-0.526***(-11.643)
Diagnostic Statistics						
Adjusted R^2	0.73	0.71	0.68	0.79	0.88	0.69
S.E. of Reg	0.087	0.09	0.095	0.077	0.059	0.093
SBC	-1.447	-1.437	-1.394	-1.764	-1.925	-1.430
DW Statistics	2.35	2.23	2.48	2.22	1.97	2.26

Note: Δ indicates differenced operator. The figure in parenthesis denotes t-statistics. ***, ** & * indicates 1%, 5 % and 10% level of significance, respectively. In F-Statistics, the figure in parenthesis indicates the p-value.

Conclusion

The study analyses investment crowd-out dynamics between public and private sector in India in a flexible accelerator theoretical setup employing ARDL cointegration estimation method controlling macroeconomic and financial variables including estimated output gap, sovereign vulnerability and inflation expectations using statistical filters and principal components. At the aggregate level, both the long and short-run coefficients support complementarity between private and public investment. Moreover, fiscal deficit, inflation expectation and sovereign vulnerability impact adversely to private capital formation, but credit flow allures. On the other hand, cost of credit appears insensitive in the long-run, though drag on private investment in short-run. Further, public investments at sectoral (infrastructure, non-infrastructure and service sectors) and industry level also complement private investment in long-run; suggestive of a homogeneous reinforcing relationship for India. However, public investment in financial services and construction crowd out private investment contemporaneously and with a year lag, respectively, in short-run. The output gap, trade openness and 1990s reform also influence private investment significantly, more at the sectoral level.

Further, one per cent increase in fiscal deficit may reduce private investment by over 50 basis points in the long-run at aggregate and sectoral level. Moreover, crowding-out bearing of fiscal deficit is significantly higher with public investment in mining and manufacturing and insignificant with infrastructure sub-sectors. Hence, public sector investment in infrastructure may tame adverse impact of expansionary fiscal policy for private investment. The sovereign vulnerability and inflation expectation dampen private investment, but credit flow lures it both in long and short-run. Also, optimizing actual output around the potential through calibrations of policy mix may be apposite to entice sustainable private investment.

Thus, fine-tuning fiscal policy with quality adjustment, containing inflation expectation and macroeconomic vulnerability would create a conducive environment for private corporate investment in India. The directed policy for infrastructure investment with adequate credit flow is desirable to support investment. The study may be extended with quarterly investment statistics at industry level allowing asymmetric and nonlinearity in standard cointegration approach.

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Appendix

Annexure Table 1. Variables Description

Variables	Description	Data Source
NCF	Net non-food credit flow from banks	DBIE, RBI
EINF	Expected inflation estimated using statistical filter techniques	Authors' estimate
GDP Def	GDP Deflator	IMF databank
GFD	Gross Fiscal Deficit	MoSPI, GoI and DBIE, RBI
INF	Consumer Price Index Inflation	DBIE, RBI
OG	Output Gap [(Actual GVA-Potential GVA)/Potential GVA*100]	MoSPI, GoI and Authors' estimate
SMV	Standard deviation of cyclic component of vulnerability measures with 3-year rolling window	DBIE, RBI and Authors' estimate
TO	Trade Openness (Export and import as per cent of GDP)	DBIE, RBI
WALR	Weighted Average Lending Rate	Basic Statistics Return, RBI
PS	GFCF in public sector	MoSPI, GoI
PRC	GFCF in private corporate sector	MoSPI, GoI
HS	GFCF in household sector	MoSPI, GoI
AFF	Public sector GFCF in Agriculture, Forestry and Fishing.	MoSPI, GoI
MINI	Public sector GFCF in Mining	MoSPI, GoI
MANU	Public sector GFCF in Manufacturing	MoSPI, GoI
EGW	Public sector GFCF in Electricity, Gas and Water supply.	MoSPI, GoI
CONS	Public sector GFCF in Construction	MoSPI, GoI
THR	Public sector GFCF in Trade, Repair, hotels & restaurants	MoSPI, GoI
TSC	Public sector GFCF in Transport , storage & communication	MoSPI, GoI
FS	Public sector GFCF in Financial services	MoSPI, GoI
RES	Public sector GFCF in Real estate, ownership of dwellings & business services	MoSPI, GoI
PADO	Public sector GFCF in Public administration and defence	MoSPI, GoI
INFR	Public sector GFCF in Infrastructure sector	MoSPI, GoI
NINFR	Public sector GFCF in non-infrastructure sector	MoSPI, GoI
CINFR	Public sector GFCF in core infrastructure sector	MoSPI, GoI
Service	Public sector GFCF in Services sector	MoSPI, GoI

Notes: DBIE: Database on Indian Economy; GDP: Gross Domestic Product; GFCF: Gross Fixed Capital Formation; GoI: Government of India; GVA: Gross Value Addition; IMF: International Monetary Fund; MoSPI: Ministry of Statistics and Programme Implementation; RBI: Reserve Bank of India

Annexure Table 2: Unit Root Test

	Augmented Dickey – Fuller (ADF)		ADF with Structural Break		Philip-Perron		Order of Integration
Variables	Level	1st Difference	Level	1st Difference	Level	1st Difference	
OG	-2.977	-6.785	-5.973		-2.709	-11.892	I(1) but I(0) with breakpoint at 2007
Ln GFD	-2.968	-6.290	-4.086	-7.669	-3.061	-6.440	I(1)
Ln NCF	-1.784	-3.547	-5.817		-1.647	-3.540	I(1) but I(0) with breakpoint at 2004
WALR	-1.537	-4.876	-6.388		-1.647	-4.893	I(1) but I(0) with breakpoint at 2013
EINF	-1.990	-3.536	-3.100	-13.829	-2.101	-3.201	I(1)
TO	-1.464	-4.871	-5.488		-1.921	-4.865	I(1) but I(0) with breakpoint at 2004
SMV	-5.285		-10.132		-10.132		I(0)
Ln PS	-1.328	-5.333	-4.264	-5.595	-1.259	-5.321	I(1)
Ln PRC	-2.377	-4.758	-3.618	-5.104	-2.114	-4.772	I(1)
Ln AFF	1.335	-6.459	-3.276	-7.466	1.335	-6.470	I(1)
Ln MINI	-3.022	-6.384	-4.111	-7.103	-3.022	-6.492	I(1)
Ln MANU	-2.414	-4.950	-5.366		-2.724	-4.926	I(1) but I(0) with breakpoint at 1998
Ln EGW	-2.598	-5.809	-4.838		-2.663	-5.817	I(1) but I(0) with breakpoint at 1998
Ln CON	-2.686	-7.823	-7.113		-2.54	-7.958	I(1) but I(0) with breakpoint at 2000
Ln THR	-2.525	-6.151	-4.191	-7.184	-2.525	-7.011	I(1)
Ln TSC	-2.047	-5.674	-3.326	-6.385	-2.023	-5.952	I(1)
Ln FS	-2.490	-7.996	-3.587	-9.646	-2.234	-15.955	I(1)
Ln RES	-0.637	-9.079	-3.817	-9.036	-1.78	-9.036	I(1)
Ln INFR*	-5.168		-6.451		-2.706	-5.938	I(1) but I(0) with breakpoint at 2000
Ln NINFR	-2.578	-5.436	-5.090		-2.749	-5.436	I(1) but I(0) with breakpoint at 1996
Ln CINFR*	-5.102		-5.792		-2.919	-6.217	I(1) but I(0) with breakpoint at 2013
Ln Service	-2.458	-5.665	-5.398		-2.496	-6.290	I(1) but I(0) with breakpoint at 1996

Note: Critical values for ADF unit root test are -4.22, -3.53 and -3.20; and -5.35, -4.86 and -4.61 for 1 %, 5 % and 10 % level of significance, without and with breakpoint, respectively. *Kwiatkowski-Phillips-Schmidt-Shin (1992) test confirm I(1) without structural break. Ln denotes natural logarithm.

Estimation of Output Gap

For estimating output gap (OG), literature uses two main variants of statistical filters for extracting the cyclical component of a series. Most of the early literature uses Hodrick-Prescott (HP) filter (Hodrick and Prescott, 1997), though the band-pass (BP) filter proposed by Baxter and King (1999) - BP(BK) got more prominence lately. Moreover, it's other variants developed by the Christiano and Fitzgerald (2003) - BP(CF), using symmetric and non-symmetric weights are also used in literature to tackle the end points problems.

We apply all the statistical filters discussed above to compute the OG by augmenting forecasted output from International Monetary Fund (IMF) to wane the end points concern of filter techniques. We observe large variations in OG estimations using different filters. The OG estimated through HP and BP(BK) are aligned, but estimation based on BP(CF) using *symmetric* and *non-symmetric* weights are highly dislocated. We also observe that all four variants of estimated OGs have positive and significant pairwise correlation coefficient. Considering pros and cons of various statistical filters, we construct principal component of output gaps for further use in econometric estimation (Figure A1).

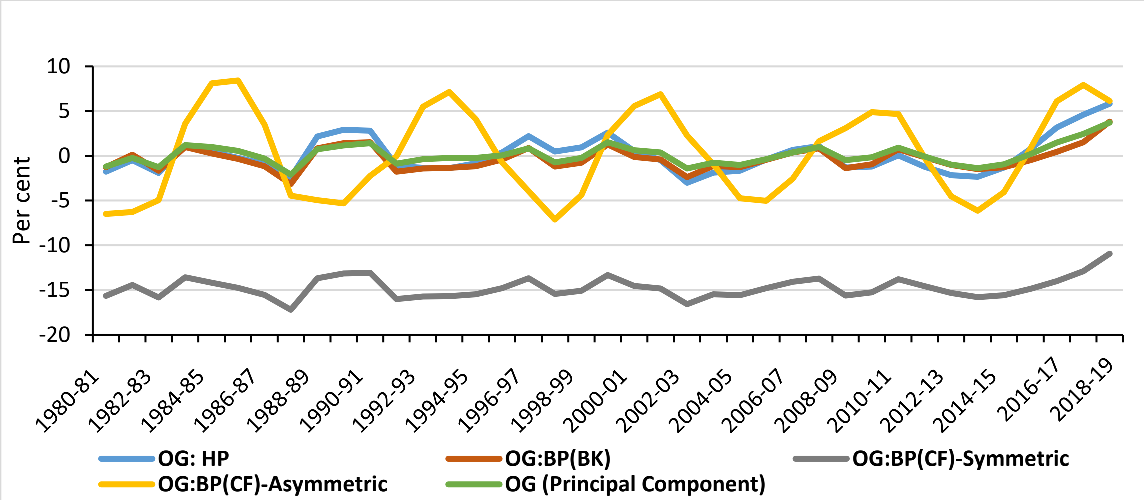

Figure A1. Estimated Output Gap

Estimation of Sovereign Macroeconomic Vulnerability (SMV)

Many studies examine the impact of macroeconomic vulnerability/volatility on output from investment point of view. The economic volatility evaluates deviation between its present value and equilibrium value and is measured by one of the three popular methods depending upon characteristics of variables and its frequency of availability namely (a) standard deviation of growth rate of a variable (b) standard deviation of the residual of an econometric regression and (c) standard deviation of the cycle isolated by a statistical filter.

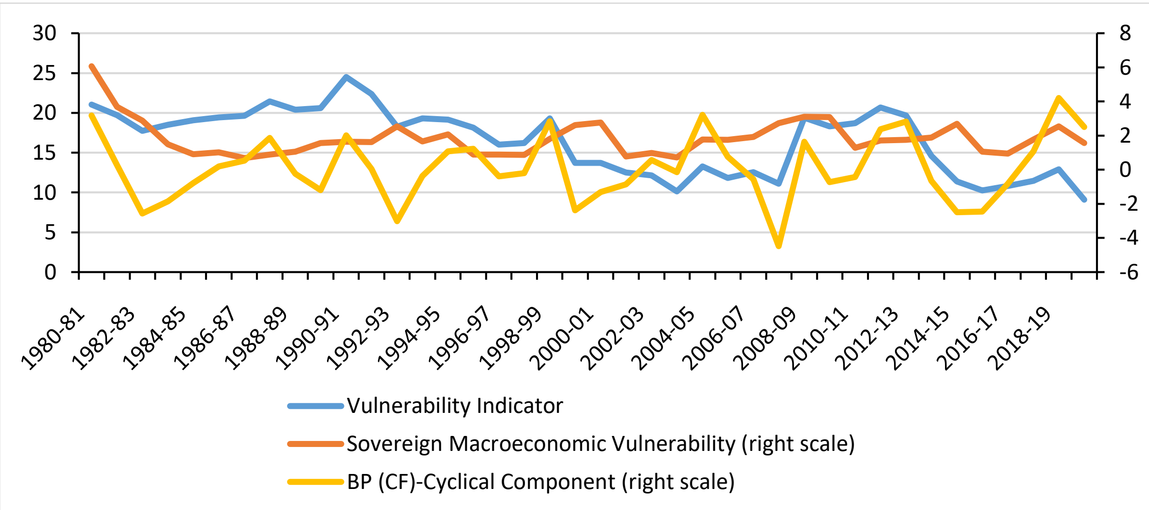

Figure A2. Vulnerability Measures

Following Becker & Mauro (2006), we compute standard deviation of estimated cyclical fluctuation of vulnerability indicator as sovereign macroeconomic vulnerability (SMV) for India, contrary to standard deviation of GDP or inflation volatility. The comprehensive measure of annual vulnerability indicator is formed combining deficits and inflation i.e., fiscal and current account deficits as ratio to GDP and GDP deflator (GoI, 2019). We employ BP(CF) with non-symmetric weight to estimate the cyclical component of the vulnerability indicator and then compute 3-year moving standard deviation considering 3-year duration of investment cycle in India as SMV (Figure A2).

Estimation of Expected Inflation

Inflation expectation measures the opinion about the future inflation trajectory. Also, the current inflation may partly derive the expectation about inflation and vice-versa. There are three primary ways to track inflation expectations: surveys of household and businesses, survey of professional forecasters, and inflation-linked financial instruments. However, the long series data on either type of surveys are not available for India. Moreover, inflation index bond/capital index bond in India was first issued in 1997, so it is impractical to get market related measure of inflation expectation. Going through the literature, low frequency components estimated using the HP filter has been considered as measure of inflation expectation (Correia, Neves, & Rebelo, 1995). We apply HP filter on CPI based inflation measure in India to estimate the trend component as measure of inflation expectation, augmenting IMF inflation forecast to tackle filter's end sample issue (Figure A3).

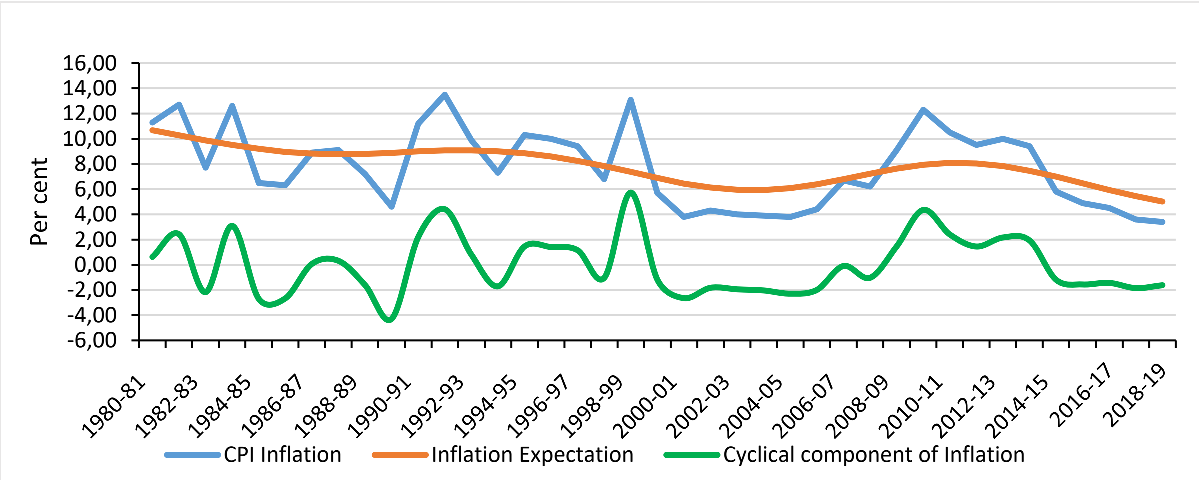

Figure A3. Inflation Decomposition

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