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King Abdullah Petroleum Studies and Research Center

Determinants of Remittance Outflows: The Case of Saudi Arabia

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March 2022 *Doi: 10.30573/KS--2022-DP05*

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Key Points

We investigate the outflow of remittances from Saudi Arabia during the period from 1970 to 2019. Many economic recessions and expansions occurred during this period. Saudi Arabia is a leading global economy and ranks first in the Gulf Cooperation Council region in terms of the outflow of remittances. Applying the cointegration and equilibrium correction methods and adjustments for small sample bias, we measure the impacts of determinants on remittances. In the long run, keeping other factors unchanged:

- a 1% increase in Saudi Arabia's gross domestic product increases the outflow of remittances by 1.7%.
- non-Saudi employment increases the remittance outflow by 1.0%.
- a 1% increase in the Kingdom's price level decreases the remittance outflow by 1.4%.
- a 1% increase in the expatriate levy reduces the remittance outflow by 0.1%.

This study's findings may be useful for macroeconomic policymaking, as the remittance outflow has numerous implications for the development of the Saudi economy. Additionally, remittances are a main channel for the leakages of money from Saudi Arabia. These leakages reduce the economic growth effects of fiscal spending multipliers.

1. Introduction

International labor migration has played a key role in the development of both advanced and developing countries. Many developing countries in Asia have relied on labor migration, mainly to the oil-rich Gulf region, to reduce both unemployment and poverty (Naseem 2007). Mansoor and Quillin (2006) explain that poverty, unemployment and low wages in developing countries are the main drivers of migration from these countries. Higher wages and the potential for improved standards of living and professional development in resource-rich countries are pull factors for migration.

Labor migration helps developing countries through various channels, such as remittances, reduced pressure on their labor markets, contact with international markets and access to technology (World Bank 2006). Developed and some developing countries face labor shortages owing to the growing demand for labor to complete large development projects. Migrant labor from developing countries can fulfill this labor demand.

For decades, migrant labor has been a key part of Saudi Arabia's economic activity. The oil boom of the 1970s created great opportunities for the Kingdom. Driven by substantial revenues from oil exports, Saudi Arabia's economy has changed dramatically over the last six decades. Since the early 1970s, Saudi Arabia has designed targeted development plans to boost economic growth and sustainable development. Private sector economic activity has been closely linked to the country's overall development path since the adoption of the first development plan (1970 to 1974). During the second development plan (1975 to 1979), a sharp increase in government spending began to shape the modern, emerging private sector. This spending has boosted industry, agriculture, health care, education, transportation and hard infrastructure in the country (Cappelen and Choudhury 2007; Al-Rushaid 2010).

During this period, high oil revenues positively impacted Saudi Arabia's economic growth and fiscal performance and strengthened its macroeconomic indicators. They laid the foundation for a prosperous economy. Migrant workers have remained an integral part of the Saudi economy since the 1970s. The share of migrants in Saudi Arabia's total population increased from 11% in 1974 to 38% in 2018 (SAMA 2020). In addition, the share of migrant workers in the Kingdom's total employment reached 76% in 2019 (GaStat 2019).

The increase in the external labor force has, in turn, led to an increase in financial flows from Saudi Arabia. The remittance outflow from Saudi Arabia grew 127 times from $267.8 million in 1972 to $33.8 billion in 2018. The outflow of remittances from Saudi Arabia was $39 billion, or 5.9% of gross domestic product (GDP), in 2015. This figure ranked second in the world after that from the United States (U.S.).

According to Naseem (2007), most migrant workers in Saudi Arabia belong to poorer classes in their home countries, and their families depend on their incomes for their livelihoods. These migrant workers immediately remit their incomes to their families. Additionally, migrant workers have limited or no investment opportunities and property/asset ownership in Saudi Arabia. These restrictions are also a major reason for their low propensity to invest in Saudi Arabia (Naufal 2011; Naufal and Genc 2012).

This substantial remittance outflow has led to concern among policymakers regarding its macroeconomic effects on the Saudi economy. Some analysts have explicitly argued that this outflow slows economic growth and negatively impacts the balance of payments (Ghosh 2006). The outflow of remittances has a negative relationship with domestic consumption and saving, and a reduction in saving, in turn, decreases investment (Al-Abri, Genc, and Naufal 2018).

Oil-based government expenditures in the Kingdom are important for its economic growth. Thus, fiscal policy plays a crucial role in Saudi Arabia's macroeconomic management, as in any other oil exporting economy (Al Moneef and Hasanov 2020; Hasanov et al. 2021). However, the outflow of remittances leads to leakages of money from the Kingdom and a low propensity for consumption among migrant workers. As a result, the public spending multiplier, which reflects the ability of public spending to stimulate economic activity, is smaller (Al-Abri, Genc, and Naufal 2018).

To reduce the dependency on foreign labor and increase employment opportunities for Saudis, the government of Saudi Arabia has shifted its policy regarding migrant workers. Since the launch of the Saudi Vision 2030 masterplan in 2016, the government has steadily taken steps to diversify its economy away from oil. In this regard, reforms and initiatives have been introduced to promote the employment of Saudi workers.

These reforms are in line with other Saudization policies, such as Nitaqat and the expatriate levy. Nitaqat is a Saudi government program to increase employment opportunities for Saudi citizens in the private sector. Under Nitaqat, the government evaluates private companies based on their nationalization performance, that is, the percentage of Saudi nationals that they employ. Certain employment sectors are restricted for foreign workers, and an expatriate levy is imposed on foreign workers and their dependents. These measures have adverse effects on migrant employment in the Kingdom and, thus, have reduced the outflow of remittances (see Figure 2).

Against this background, this study aims to assess the impacts of the determinants on the outflow of remittances from Saudi Arabia and propose policy insights. Additionally, understanding the effects of measures introduced by Saudi authorities, such as the expatriate levy, on remittance outflow is important. We applied a cointegration and equilibrium correction methodology to the data to assess the long- and short-run impacts of the key driving factors of the outflow of remittances.

We find that GDP, non-Saudi employment, the domestic price level and the expatriate levy have statistically significant long- and short-run impacts on the outflow of remittances. We also find that 20.4% of the previous year's deviations of outflow remittances from its long-run path, caused by interventions including policy shocks, adjusts back to that path in the present year.

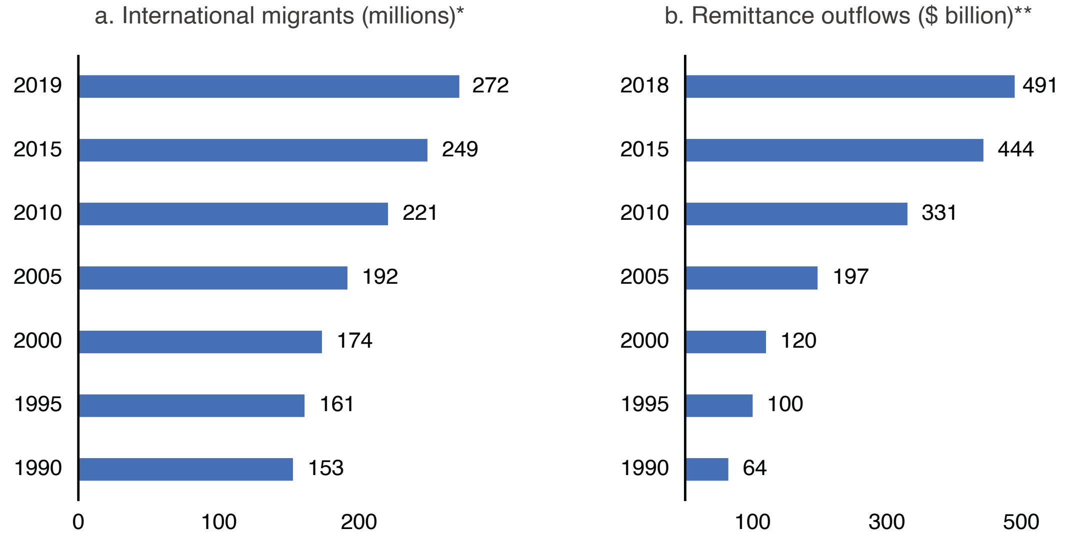
This study contributes to the existing literature in four ways. First, Saudi Arabia is one of the largest destinations for migrant workers. It may be the largest source of migrant remittances among the Gulf Cooperation Council (GCC) countries (see Table 1). Thus, Saudi Arabia is an important case study for understanding the determinants of remittance outflows. However, existing studies largely focus on the inflow of remittances and their significance for the economic development of the receiving countries; academic researchers and policymakers have focused less on the outflow of remittances.[1] Some studies (e.g., Rahman, Abdel-Mahmoud [2006]; Alkhathlan [2013]; Termos, Naufal, and Genc [2013]; Al-Haddad and Choukir [2015]; Hathroubi and Aloui [2016]; Naufal and Genc [2017]; Al-Abri, Genc, and Naufal [2018]) have investigated the role of the outflow of remittances in Saudi Arabia. However, they focused on the macroeconomic effects of these remittances, such as their effects on economic growth, monetary policy, and inflation. Very few studies investigate the determinants of remittance outflows from Saudi Arabia. In this regard, this is one of the pioneer studies for Saudi Arabia as it investigates the outflow of remittances by addressing integration and cointegration properties of the time series data, while accounting for small sample biases to elicit grounded insights for policymakers. Second, no previous studies have covered the recent low oil price environment coupled with the unprecedented fiscal and energy price reforms, which have considerable implications for the remittance outflows.

The rest of this paper is structured as follows. In Section 2, we provide an overview of migration to Saudi Arabia and trends in the remittance outflows. We also compare Saudi Arabia with the rest of the world. Section 3 surveys existing studies of remittances from Saudi Arabia. In Section 4, we describe our theoretical framework, and in Section 5, we present the data sources and definitions of variables. The methodological approach used for the analysis is described in Section 6. In Section 7, we present the estimation and testing results, and in Section 8, we discuss the empirical findings. Lastly, Section 9 concludes the study with some policy insights.

2. Remittance Outflows From Saudi Arabia: Some Stylized Facts

In the last three decades, the international migrant population has increased. The number of international migrants worldwide increased from 153 million in 1990 to 272 million in 2019. This growth corresponds to an average increase of 2.7% per year (see Figure 1a) (U.N. 2019). In other words, the number of international migrants increased by 1.8 times in the last three decades. The annual average rate of change in international migration was 1.6% between 1990 and 2005 and 2.8% between 2005 and 2019.

Figure 1. Worldwide stock of international migrants and remittance outflows.



Source: *United Nations (2019). **World Bank (2020a).

Statistics from the World Bank (2020a) show that the flow of remittances increased sharply from $64.3 billion in 1990 to $491 billion in 2018. Remittances from migrants were 7.6 times greater in 2018 than in 1990 (see Figure 1b).

Worldwide, Saudi Arabia is an important destination for migrant workers and a major source of remittances after the U.S. Averages for 1980 to 2019 show that in absolute and percentage terms, the U.S. and Saudi Arabia had the largest outflows of remittances (see Table 1). Other important sources of remittances were the United Arab Emirates, Switzerland and Germany.

The number of migrant workers in Saudi Arabia increased from 5 million in 1990 to 13 million in 2019, corresponding to an average annual increase of 8.1%. In 2019, migrant workers were 38.3% of Saudi Arabia's total population, whereas they were 30.8% of the total population in 1990. Saudi Arabia is also the second-largest destination of migrant workers in terms of the percentage of the population (see Table 1).

2. Remittance Outflows From Saudi Arabia: Some Stylized Facts

Table 1. Top five countries in terms of migrants and the outflow of remittances.

10-year average	U.S.	UAE	KSA	SWIT	GER	U.S.	UAE	KSA	SWIT	GER
	Remittance outflows ($ billion)[a]					Remittance outflows (% of GDP)				
1980-1989	6.6		5.5	3.6	4.9	0.2		5.0	2.4	0.5
1990-1999	21.3		14.8	8.6	8.8	0.3	4.4	10.4	3.0	0.4
2000-2009	46.1	6.5	17.9	11.2	12.4	0.4	3.4	5.9	2.6	0.4
2010-2019	60.2	31.4	34.2	25.3	19.9	0.3	8.1	4.8	3.7	0.5
Year	International migrant stock (millions)[b]					Migrant stock (% of total population)				
1990	23.3	1.3	5.0	1.4	5.9	9.2	71.5	30.8	20.9	7.5
2000	34.8	2.4	5.3	1.6	9.0	12.4	78.1	25.5	22.0	11.0
2010	44.2	7.3	8.4	2.1	9.8	14.3	85.6	30.7	26.6	12.1
2019	50.7	8.6	13.1	2.6	13.1	15.4	87.9	38.3	29.9	15.7

Source: a. Data on remittance outflows and GDP are taken from World Bank (2020b). b. Data on the migrant stock and total population are taken from U.N. (2019).

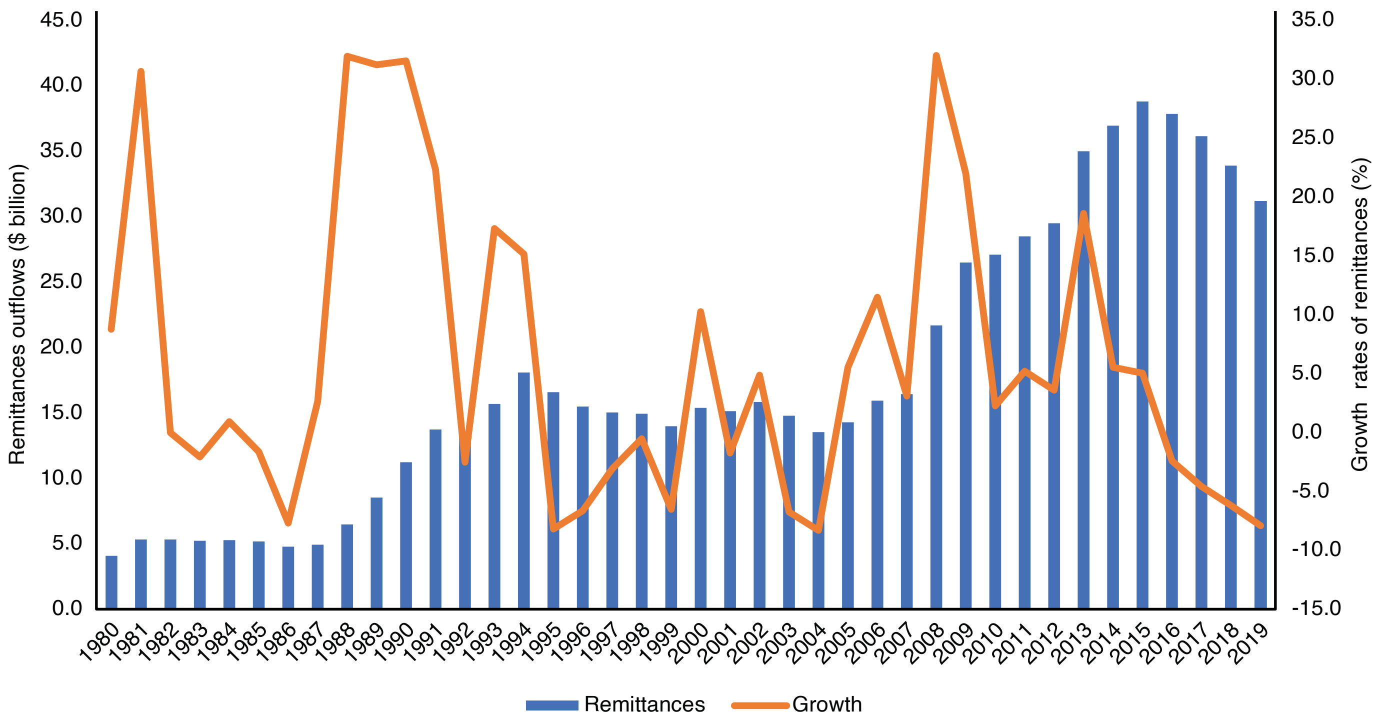
Notes: UAE = United Arab Emirates; KSA = Saudi Arabia; SWIT = Switzerland; GER = Germany.

Table 1 shows that remittance outflows from Saudi Arabia as a percent of GDP varies substantially compared with other top remittance-sending countries. This value increased gradually from 2.5% of GDP in 1980 to a peak of 13.4% of GDP in 1994. However, after 1994, remittance outflows as a percent of GDP declined, with no systematic pattern. In the 2000s and 2010s, remittance outflows as a percent of GDP were 6% and 4.8%, respectively. Overall, the annual average remittance outflow from Saudi Arabia between 1980 and 2019 was 6.5% of GDP. The corresponding percentages for the U.S., Switzerland and Germany were 0.3%, 2.9% and 0.5%, respectively.

In absolute terms, the outflow of remittances from Saudi Arabia has continuously increased over the last five decades, except during the last three years. Remittance outflows from Saudi Arabia grew more than ninefold between 1980 and 2015, with average annual growth of 7.2%. Figure 2 similarly shows that, since 1980, the outflow of remittances (blue bar chart) from Saudi Arabia has trended upward overall, with high volatility.

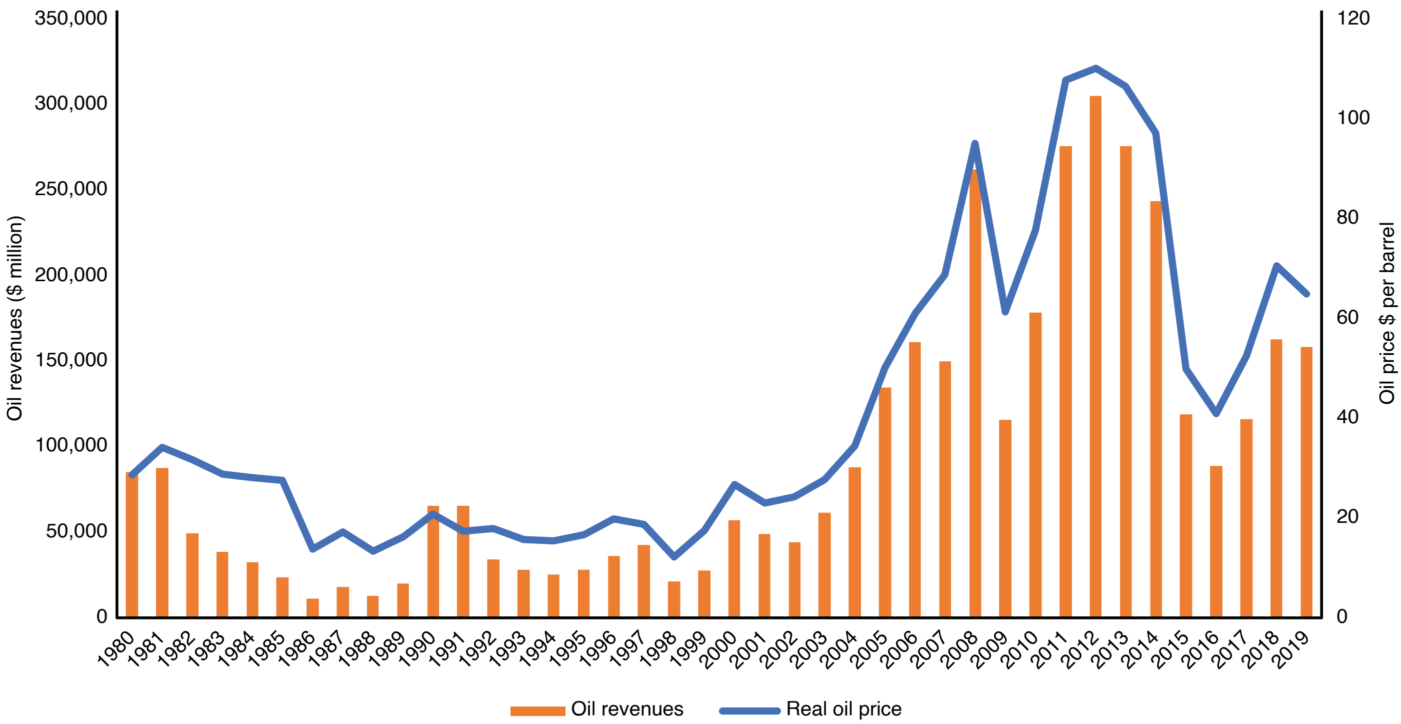
Figure 2 illustrates two periods in which remittance outflows rapidly increased: 1987 to 1994 and 2006 to 2015. The outflow of remittances from Saudi Arabia increased during the oil boom years of the 1970s and early 1980s. However, it declined in the mid-1980s as oil prices fell, the budget deficit widened, and the government set limits on hiring foreign workers (Ratha 2005). Similarly, remittance outflows from Saudi Arabia fell substantially after 2015. This fall was due to the oil price collapse and the subsequent reduction in Saudi Arabia's oil revenues, as illustrated in Figure 3. Nationalization policies and the expatriate levy may have also driven the gradual decline in remittance outflows from Saudi Arabia.

Figure 2. The level and growth rate of remittance outflows from Saudi Arabia (1980-2019).



Source: World Bank (2020a).

Figure 3. Trend in the oil price and oil revenues (1980-2019).



Source: Saudi Arabia Monetary Agency, Annual Report 2020.

3. Literature Review

Two strands of literature describe the impacts of immigration on a host country's economy. One strand argues that immigration negatively affects the host economy through labor market distortions; remittance outflows; and the income distribution. For example, Alkhathlan (2013) argues that the outflow of remittances can be considered as a withdrawal of money from circular flows. This withdrawal reduces economic activity and negatively affects the host country's aggregate demand in the short run, but it has no long-run effects. The other strand of literature stresses the positive role of migrant workers in the host economy. These workers can address labor shortages and the problems associated with an aging population (e.g., Hathroubi and Aloui [2016]). Remittances can also have positive effects in remittance-sending countries by reducing inflation (Termos, Naufal, and Genc 2013).

Several previous studies analyze remittances from Saudi Arabia. Abdel-Rahman (2006) investigates the determinants of the outflow of remittances from Saudi Arabia over the period from 1975 to 2001. The study uses several indicators to identify the determinants of the remittance outflows from Saudi Arabia. These indicators include GDP per capita, wages per worker, return and parity conditions, composite indices related to socioeconomic factors and risk indicators. The study concludes that GDP per capita and wages have significant positive impacts on the outflow of remittances per employee. The nominal interest rate and the differential parity between the host country and the country of origin have significant negative impacts on the outflow of remittances. The Granger causality test results indicate bidirectional causality between the per worker remittance outflows and per capita output. However, Abdel-Rahman (2006) ignores the time series properties of the data, such as non-stationarity and cointegration. The literature shows that ignoring these properties can lead to spurious findings and, thus, misleading policy recommendations (e.g., Yule [1926]; Engle and Granger [1987]; Banerjee et al. [1993]; Enders [2015]).

Hathroubi and Aloui (2016) analyze the dynamic relationship between the outflow of remittances and macroeconomic indicators in Saudi Arabia for the period from 1980 to 2013. They use the Granger causality test, the wavelet power spectrum, the cross-spectrum wavelet and the coherence wavelet. The results show that remittance outflows from Saudi Arabia are strongly correlated with the real growth rate, employment and government expenditures. In addition, the results show that the positive link between remittance outflows and real economic growth is more evident in the short run.

Alkhathlan (2013) investigates the impact of remittance outflows on economic growth in Saudi Arabia from 1970 to 2010. He uses the autoregressive distributed lag (ARDL) and equilibrium correction model (ECM) approaches. The results show that remittance outflows weaken economic growth in the short term and have no impact in the long term.

Rahmouni and Debbiche (2017) examine the effect of remittance outflows on economic growth for Saudi Arabia from 1970 to 2014. They use the ARDL and ECM approaches to estimate the short- and long-run impact of remittance outflows on economic growth. The findings show that remittance outflows from Saudi Arabia have no significant impact on economic growth, both in the short and long run. According to the authors, the decreasing share of remittance outflows in GDP, and the negative effect of remittance outflows on inflation are the possible reasons for the neutral effect of remittance outflows on economic growth in Saudi Arabia.

Salameh (2019) analyzes the impact of remittance outflows on economic growth in Saudi Arabia from 1972-2016 by using the Bai-Perron test and ordinary least squares (OLS) approaches. The results reveals that remittance outflows have a negative and statistically significant impact on economic growth.

Haddad and Choukir (2015) examine the time-varying impulse responses of non-oil GDP, investment and the current account balance to structural remittance outflow shocks. They use a time-varying parameters vector autoregressive model and data from Saudi Arabia for the period from 1970 to 2012. The study reveals that the responses depend on the magnitude of the structural volatility of remittance outflows. Highly volatile remittances seem to have had persistent negative effects on non-oil GDP, investment and the current account balance in the 1970s and 1980s. Moreover, the results indicate that the time-varying response of non-oil GDP to remittance shocks is negative from 1980 to 1992 and positive otherwise.

Al-Abri, Genc, and Naufal (2018) examine the macroeconomic impact of public spending on output growth with and without considering the role of remittances. They use the vector autoregressive (VAR) approach and data spanning the period from 1980 to 2015. Their findings show that remittance outflows from Saudi Arabia have a weak impact on public spending and, thus, on output growth in Saudi Arabia.

Some other previous studies investigate remittance outflows from the GCC more broadly. Naufal and Termos (2009) investigate the impacts of changes in the crude oil price, government spending and global GDP on the outflow of remittances. They use unbalanced panel data from six GCC countries for 1971 to 2004 and apply the ordinary least squares (OLS) and fixed effects (FE) approaches. The results show that outflow of remittances from GCC countries responds positively and significantly to the oil price but is inelastic. However, they do not consider the economic activity of the host countries to be a determinant of the outflow of remittances. A seminal study by Swamy (1981) shows that the change in economic activity in the host country is an important determinant of remittance outflows.

Termos, Naufal, and Genc (2013) use panel data from the GCC countries for the period from 1972 to 2010. They analyze the effects of remittance outflows on inflation, also using the OLS and FE approaches. The results show that remittance outflows exert deflationary pressure on the remitting countries. The overwhelming volume of remittances can be seen as a tacit monetary stabilization tool for GCC countries.

The aforementioned studies analyze the impact of remittance outflows on macroeconomic indicators. None of the studies have focused on the determinants of remittance outflows, except Abdel-Rahman (2006). Abdel-Rahman (2006) focuses on the determinants of remittance outflows from Saudi Arabia. However, he ignores the time series properties of the data, such as non-stationarity and cointegration. Non-compliance with the stationarity and cointegration assumption in the analysis of time series variables can lead to an incorrect and unreliable statistical conclusion. Additionally, Abdel-Rahman (2006) has not used data from recent years. The low oil prices and reforms to fiscal policy and domestic energy prices in recent years may have had considerable implications for remittance outflows. In this study, we fill this gap in the literature on remittance outflows from Saudi Arabia by addressing these points.

4. Theoretical Framework for the Remittance Outflow

4.1. Labor Demand

We assume that output (Y_t^d) in the host country (i.e., the labor-receiving country) is produced according to a Cobb-Douglas production function. The function includes three production factors: physical capital (K), the domestic labor supply (L_t^d) and foreign labor (L_t^f).

$$Y_t^d = AK_t^{\alpha} \left(L_t^d\right)^{\beta} \left(L_t^f\right)^{1-\alpha-\beta}, \qquad (1)$$

where α, β and 1-α- β are the output elasticities of capital, domestic labor and foreign labor, respectively.

In a competitive economy with constant returns to scale, factors of production are paid according to their marginal products. Thus, the marginal product of foreign labor (L^f) is equal to the wage of foreign workers (W^f).

$$\frac{(1-\alpha-\beta)Y_t^d}{L_t^f} = W_t^f \quad \rightarrow \quad W_t^f L_t^f = (1-\alpha-\beta)Y_t^d \qquad (2)$$

Equation (2) represents foreign workers' earnings, which they can remit to their home country. We assume that foreign workers can transfer a portion of their total income, minus the cost of living and expatriate tax, to their home country. Thus, the equation of remittance (REM) can be written as follows:

$$REM_t = W_t^f L_t^f e^{-(lnP_t^{\vartheta} + lnEXP_t^{\tau})}. \qquad (3)$$

In Equation (3), P_t and EXP_t represent the cost of living for foreign workers in Saudi Arabia and the expatriate levy, respectively. ϑ and τ are the parameters of P_t and EXP_t, respectively.

We assume that 0 <ϑ<1 and 0<τ<1. These restrictions indicate that foreign workers in Saudi Arabia spend a fraction of their total income on living costs and the expatriate levy.

One can argue against including the expat levy variable in equation (3), i.e., in the long-run model, as it was implemented in 2017. However, the following three reasons encourage us to include it in the theoretical model. First, if we exclude it from the theoretical model and the long-run estimates in section 7, it almost makes no change to the coefficients of the other explanatory variables (the results are available on request). Second, it is a policy variable, like tax in a conventional economy, and therefore it is useful for the policy-making process to quantify its impact in both the long and short run. Third, it has values for three years (2017-2019) and sometimes a period of more than two years is considered medium to long term in macroeconomic studies, especially in fiscal multiplier analyses (see, for example, Espinoza and Sendjahi [2011]).

We take the natural log of Equation (3), as follows:

$$\ln REM_t = \ln W_t^f + lnL_t^f - \vartheta lnP_t - \tau lnEXP_t. \quad (4)$$

Taking the log of Equation (2) gives the following equation:

$$\ln\left(L_t^f\right) = \ln(1-\alpha-\beta) + \ln\left(Y_t^d\right) - \ln\left(W_t^f\right). \quad (5)$$

4.2. Labor Supply

In a country with a labor surplus, the total available labor force can be split into two components. These components are the labor employed in the domestic market (L^h) and the labor supplied to the foreign market (L^f). Like Yoshino, Taghizadeh-Hesary, and Otsuka (2020), we use the following utility function:

$$U\left(C_t^h, L_t^h, L_t^f\right) = \frac{\left(C_t^h\right)^{1-\delta}-1}{1-\delta} - \frac{\left(L_t^h\right)^{1-\sigma}-1}{1-\sigma} - \frac{\left(L_t^f\right)^{1-\rho}-1}{1-\rho}. \quad (6)$$

U_t is the utility function of the migrant-sending country. This function consists of the consumption of goods (C_i) and the supply of labor to the domestic and foreign markets. The parameter δ is the elasticity of household consumption. σ and ρ are the elasticity of the labor supply in the domestic and foreign labor markets, respectively. The utility function depends positively on consumption (C_i) and negatively on the amount of labor supplied.

The budget constraint can be written as follows:

$$C_t^h + S_t = W_t^h L_t^h + W_t^f L_t^f \, e^{-(lnP_t^{\vartheta} + lnEXPL_t^{\tau})}. \quad (7)$$

For utility maximization, we set a Lagrange function:

$$\Gamma = U\left(C_t^h, L_t^h, L_t^f\right) - \lambda\left(C_t^h + S_t - W_t^h L_t^h - W_t^f L_t^f e^{-(lnP_t^{\vartheta} + lnEXPL_t^{\tau})}\right). \quad (8)$$

The first-order conditions with respect to consumption and the labor supply in the domestic and foreign markets are:

$$\frac{\partial\Gamma}{\partial C_t^h} = \left(C_t^h\right)^{-\delta} - \lambda = 0 \rightarrow \quad \left(C_t^h\right)^{-\delta} = \lambda \quad (9)$$

$$\frac{\partial\Gamma}{\partial L_t^h} = -\left(L_t^h\right)^{-\sigma} + \lambda W_t^h = 0 \rightarrow \quad \left(L_t^h\right)^{-\sigma} = \lambda W_t^h \quad (10)$$

$$\frac{\partial\Gamma}{\partial L_t^f} = -\left(L_t^f\right)^{-\rho} + \lambda W_t^f e^{-(lnP_t^{\vartheta} + lnEXP_t^{\tau})} = 0 \rightarrow$$

$$\left(L_t^f\right)^{-\rho} = \lambda\left(W_t^f e^{-(lnP_t^{\vartheta} + lnEXPL_t^{\tau})}\right) \quad (11)$$

$$\frac{\partial\Gamma}{\partial \lambda_t} = \left(C_t^h + S_t - W_t^h L_t^h - W_t^f L_t^f e^{-(lnP_t^{\vartheta} + lnEXPL_t^{\tau})}\right) = 0. \quad (12)$$

Equation (11) can be written in log linearization form as:

$$-\rho \ln L_t^f = ln\lambda + \ln W_t^f - \vartheta lnP_t - \tau lnEXPL_t \quad (13)$$

$$\ln L_t^f = \frac{ln\lambda}{\rho} - \frac{1}{\rho}\ln W_t^f + \frac{\vartheta}{\rho} lnP_t + \frac{\tau}{\rho} lnEXP_t. \quad (13a)$$

From Equation (5) and (13a), it follows that:

$$\frac{ln\lambda}{\rho} - \frac{1}{\rho}\ln W_t^f + \frac{\vartheta}{\rho} lnP_t + \frac{\tau}{\rho} lnEXPL_t = \ln(1-\alpha-\beta) + lnY_t^d - lnW_t^f$$

$$\left(1-\frac{1}{\rho}\right)\ln W_t^f = \ln(1-\alpha-\beta) - \frac{ln\lambda}{\rho} - \frac{\vartheta}{\rho} lnP_t - \frac{\tau}{\rho} lnEXPL_t + lnY_t^d$$

$$\ln W_t^f = \frac{\rho}{\rho-1}\ln(1-\alpha-\beta) - \frac{ln\lambda}{\rho-1} - \frac{\vartheta}{\rho-1} lnP_t - \frac{\tau}{\rho-1} lnEXPL_t + \frac{\rho}{\rho-1} lnY_t^d. \quad (14)$$

By substituting lnW_t^f from Equation (14) into Equation (4), we obtain Equation (15), which shows the determinants of remittances.

$$\mathrm{Ln}\, REM_t = \frac{\rho}{\rho-1}\ln(1-\alpha-\beta) - \frac{ln\lambda}{\rho-1} - \frac{\vartheta}{\rho-1} lnP_t - \frac{\tau}{\rho-1} lnEXPL_t + \frac{\rho}{\rho-1} lnY_t^d + lnL_t^f - \vartheta lnP_t - \tau lnEXPL_t$$

$$\ln REM_t = \frac{\rho}{\rho-1}\ln(1-\alpha-\beta) - \frac{ln\lambda}{\rho-1} - \frac{\rho\vartheta}{\rho-1} lnP_t - \frac{\rho\tau}{\rho-1} lnEXPL_t + \frac{\rho}{\rho-1} lnY_t^d + lnL_t^f. \quad (15)$$

Equation (15) can be written as

$$rem_t = \varphi_0 + \varphi_1 p_t + \varphi_2 expl_t + \varphi_3 y_t^d + \varphi_4 l_t^f + \varepsilon_t \quad (16)$$

$$\varphi_0 = \frac{\rho}{\rho-1}\ln(1-\alpha-\beta) - \frac{ln\lambda}{\rho-1}, \quad \varphi_1 = -\frac{\rho\vartheta}{\rho-1}, \quad \varphi_2 = -\frac{\rho\tau}{\rho-1}, \quad \varphi_3 = \frac{\rho}{\rho-1},$$

where t represents time and rem is the outflow of remittances from Saudi Arabia. y_t^d is GDP, l_t^f is non-Saudi employment and p is the price level in the Kingdom, measured by the GDP deflator. We expect y_t^d and, l_t^f to have positive impacts on the remittance outflow, meaning that the expected signs of the coefficients φ_2 and φ_3 are positive. The price level in the Kingdom is expected to lower the outflow of remittances, and, thus, the expected sign of φ_1 is negative.

5. Data

We use annual data over the period from 1970 to 2019. Table 2 provides a detailed description of each variable and its data source.

Table 2. Variables and their descriptions.

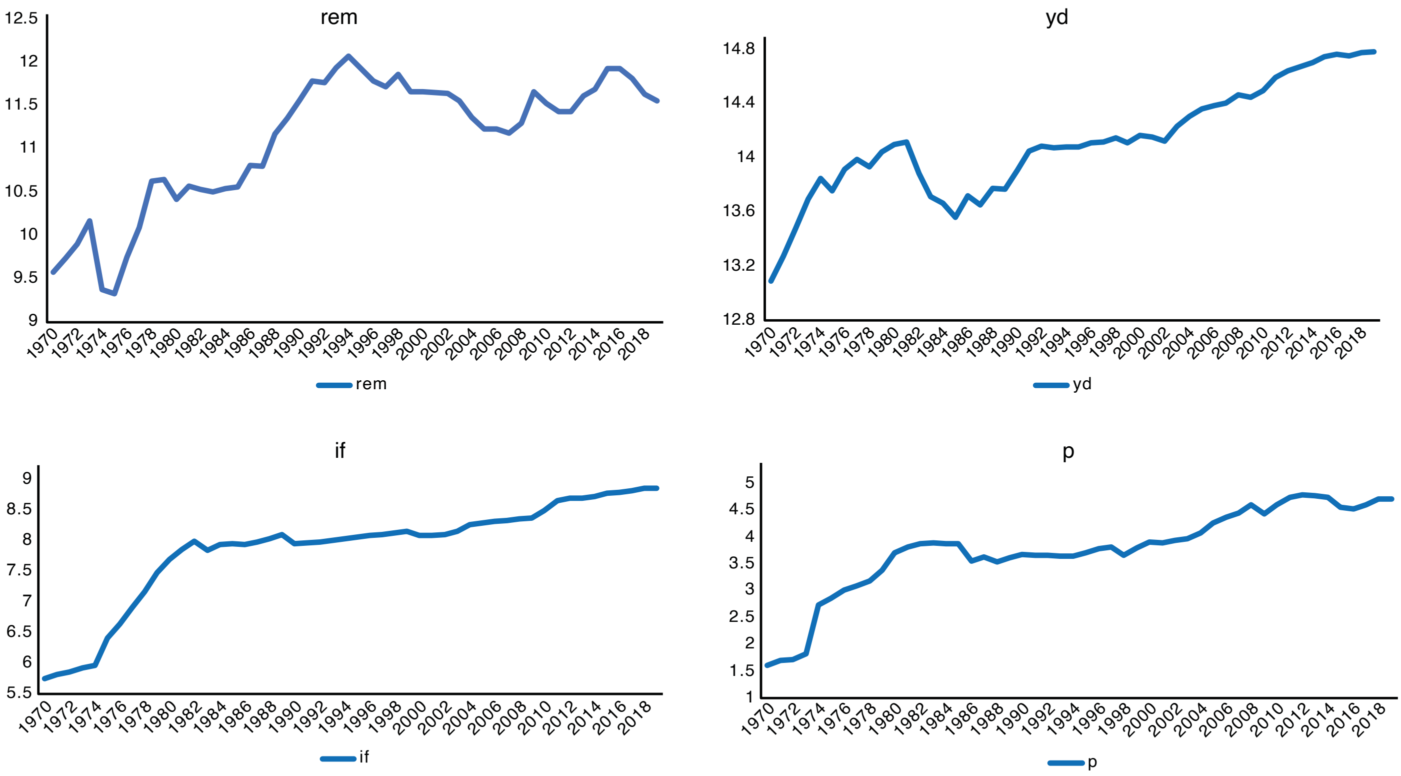
Notation	Name	Definition and source
REM	Remittance outflow, real, in million riyals at 2010 prices.	Data for remittance outflows in U.S. dollars is taken from the World Bank (2020b). It is multiplied by the bilateral nominal exchange rate of Saudi riyals to U.S. dollars to place it in terms of millions of riyals. Lastly, the resulting series is deflated by the price level (i.e., the GDP deflator, 2010=100) to convert it to real values.
Y^d	GDP, real, in million riyals at 2010 prices.	The data are obtained from SAMA (2020).
L^f	Total non-Saudi employment, in thousands of people.	Data through 2016 are retrieved from the former Central Department of Statistics and Information, currently known as the General Authority for Statistics. The values for 2017 and 2018 are obtained from Oxford Economics Global Economic Model database.
P	GDP deflator, 2010=100.	The variable is calculated as the ratio of real GDP in million riyals at 2010 prices to nominal GDP in million riyals at current prices. Nominal GDP values are also from SAMA (2020).[2]
$EXPL$	Expatriate levy, real, in million riyals at 2010 prices.	The nominal values of this variable are calculated by Hasanov et al. (2020) using information from the Fiscal Balance Program (FBP 2017) and other sources. Then, the nominal values are deflated by the price level (i.e., the GDP deflator, 2010=100) to convert them into real values.

Source: The authors calculation based on the data sources.

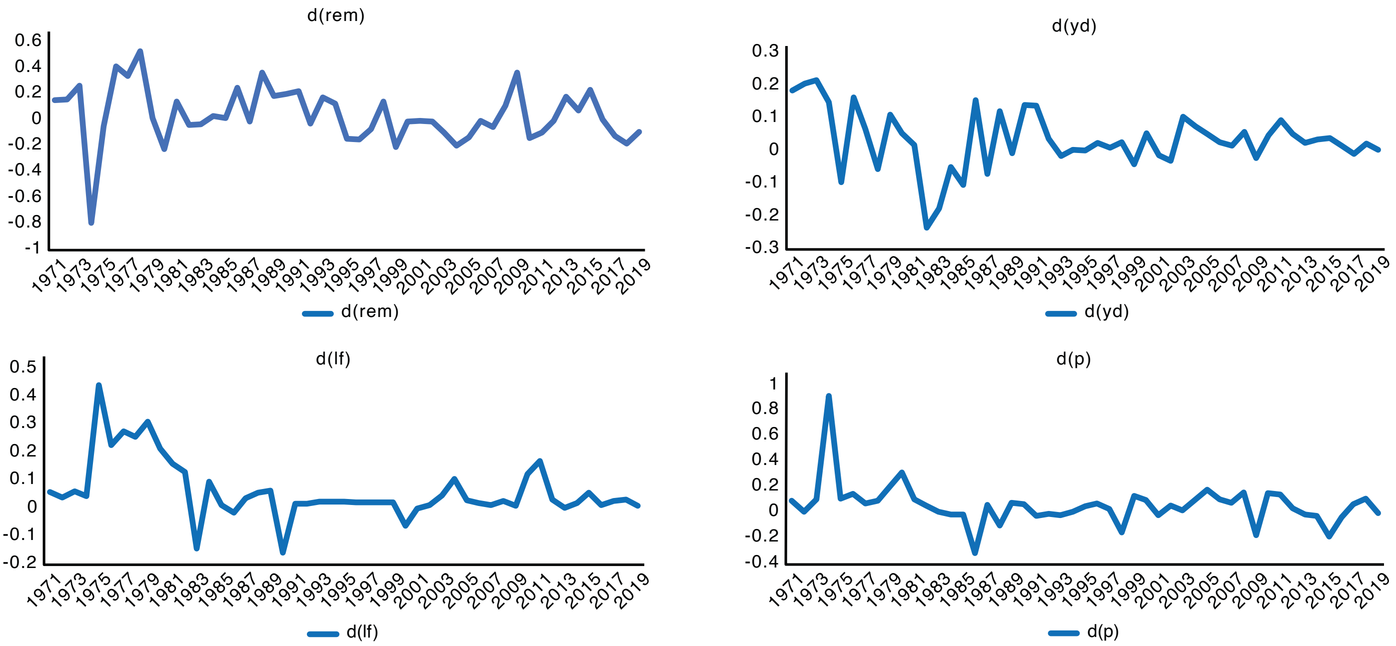
Panels A and B in Figure 4 illustrate the natural logarithmic transformations (denoted by the lower-case letters) and growth rates of the variables documented in Table 2.[3]

Figure 4. Graphs of the logarithmic levels and growth rates of the variables.

Panel A. Logarithms of the variables



Panel B. Growth rates of the variables



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6.1. Augmented Dickey-Fuller Unit Root Tests

Since the variables are time-series variables, they must be tested for unit root properties before performing empirical analysis. We use the augmented Dickey-Fuller (ADF) unit root test to check the integration orders of the variables. The general form of the ADF unit root test for a given variable y can be written as follows:

$$\mathrm{d}(y_t) = \beta_0 + \beta_1 t + \beta_2 y_{t-1} + \sum_{i=1}^{p} \gamma_i d(y_{t-i}) + \varepsilon_t. \quad (17)$$

In equation (17), β_0 is the intercept, β_1 is the coefficient of the time trend and t is the linear time trend. D is the difference operator. Lagged values of the dependent variable are included on the right side of the equation to correct for autocorrelation. p is the maximum lag order of the autoregressive process.

The unit root test is then carried out under the null hypothesis of $\beta_2=0$. The alternative hypothesis is $\beta_2<0$. If the t-statistic of β_2 is greater than the corresponding ADF critical value in absolute terms, we can reject the null hypothesis of a unit root.

Macroeconomic time series mostly exhibit stochastic trend behavior, but sometimes a deterministic trend is characteristic for a given time series. If the deterministic trend is a significant part of the data-generating process, its omission from the unit root test may produce biased results. Thus, a linear time trend is included in the unit root tests for the variables in case they exhibit any statistically significant deterministic trending behavior.

6.2. Cointegration Analysis

Once the integration order of the variables is identified, we test for the existence of a cointegrating relationship. By definition, n variables can establish at most $n-1$ cointegrating relationships (e.g., Engle and Granger [1987]). However, only the system-based cointegration tests, such as Johansen's reduced rank approach, can determine the number of cointegrating relationships (Johansen 1988; Johansen and Juselius 1990, 1992). Other cointegration tests, such as the Engle-Granger (Engle and Granger 1987) and ARDL bounds tests (Pesaran and Shin 1998; Pesaran, Shin, and Smith 2001), can only reveal whether variables are cointegrated or not.

Incorrectly determining the number of cointegrating relationships among the variables can lead to information loss and, more importantly, to omitted variable bias if the equilibrium correction term of the other cointegrating relationship is statistically significant in the short-run specification of the variable of interest (e.g., see Dibooglu and Enders [1995]; Badinger [2004]; Enders [2015]). If the Johansen method concludes that not more than one cointegrating relationship exists among the variables, then we can apply the ARDL bounds test as a robustness check (Pesaran and Shin 1998; Pesaran, Shin, and Smith 2001). We mainly use the ARDL method to estimate the level relationship between variables with one cointegrating relationship, as this method outperforms other methods in small samples (e.g., see Pesaran and Shin [1998]). For both tests, we apply a small sample bias correction for further robustness – we follow Reinsel and Ahn (1992) and Reimers (1992) for the Johansen test and Narayan (2005) for the bounds test.

Once the coefficients of the long-run relationship are estimated, we conduct a short-run analysis. We use an ECM to analyze the short-run relationships among the variables. We estimate the ECM in the general-to-specific modeling strategy framework (e.g., see Campos, Ericsson, and Hendry [2005]). In the remainder of this section, we describe the Johansen and ARDL methods in more detail.

6.2.1. Johansen Cointegration Method

The vector error correction model (VECM), developed by Johansen (1988) and Johansen and Juselius (1990), can be expressed as follows.

$$\Delta y_t = \Pi y_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \, \Delta y_{t-i} + \mu + \varepsilon_t \qquad (18)$$

where y_t is an $n \times 1$ vector of the n endogenous (i.e., modeled) variables of interest. μ is an $n \times 1$ vector of constants, and Γ is an $n \times (k-1)$ matrix of short-run coefficients. ε_{tis} an $n \times 1$ vector of white noise residuals, and Π is an $n \times n$ matrix of coefficients.

If the matrix Π has reduced rank, meaning that $0 < r < n$, then it can be divided into two elements. These elements are an $n \times r$ matrix of loading coefficients α and an $n \times r$ matrix of cointegrating vectors β. α represents the importance of the cointegration relationships in the system's individual equations and the speed of adjustment to disequilibrium. β indicates the long-term equilibrium relationship. It follows that $\Pi=\alpha\beta'$.

When testing for cointegration using Johansen's reduced rank regression approach, the following logic applies. First, the process involves estimating the matrix Π in an unrestricted form. Second, it involves testing whether the restriction implied by the reduced rank of Π can be rejected. In other words, the rank of Π is determined by the number of its characteristic roots that are different from zero. This rank characterizes the number of independent cointegrating vectors.

To conduct this analysis, we first estimate an unrestricted VAR with the maximum lag order. Then, we estimate a restricted VAR with the optimal lag order, with Gaussian errors and the stability specified. Finally, we transform the restricted VAR into a VECM and perform the cointegration test.

6.2.2 ARDL Model

This study also uses the ARDL approach to cointegration, proposed by Pesaran, Shin, and Smith (2001). This approach has several advantages. It can be applied irrespective of whether the regressors are I(0), I(1) or a mixture of the two. It captures both short-run and long-run dynamics when testing for the existence of cointegration. It offers explicit tests for the existence of a unique cointegration vector rather than assuming that it exists. It is preferable in small samples, as Pesaran and Shin (1998) show.

The ARDL specification of equation (1) can be written as follows:

$$\Delta rem_t = \beta_0 + \beta_1 rem_{t-1} + \beta_2 \, y^d_{t-1} + \beta_3 l^f_{t-1} + \beta_4 p_{t-1} + \beta_5 expl_{t-1} + \sum_{i=1}^{p0} \gamma_i \Delta rem_{t-i} + \sum_{i=0}^{p1} \delta_i \Delta y^d_{t-1} + \sum_{i=0}^{p2} \theta_i \Delta \, l^f_{t-1} + \sum_{i=0}^{p3} \vartheta_i \Delta p_{t-i} + \sum_{i=0}^{p4} \pi_i \Delta expl_{t-i} + \epsilon_t . \qquad (19)$$

As in the Johansen method, we first estimate Equation (19) with the maximum lag order of the variables. Then, we determine the optimal lag orders ($p0, p1, p2, p3$ and $p4$) for the variables. One advantage of the ARDL method is that it can provide different optimal lag orders for different variables. When we estimate Equation (19) with the optimal lag orders for the variables, we apply an F-bounds test. This test can determine the joint significance of the lagged level variables. The null hypothesis of no cointegration in Equation (19) is $H_0 : \beta_1=\beta_2=\beta_3= \beta_4= \beta_5=0$, and the alternative hypothesis is $H_1 : \beta_1\neq\beta_2\neq\beta_3\neq \beta_4\neq \beta_5\neq 0$.

The bounds test provides two asymptotic critical values. The lower critical value assumes that the explanatory variables are stationary in levels (i.e., $I(0)$). The upper critical value assumes that the explanatory variables are non-stationary in levels but are stationary in first differences (i.e., $I(1)$). If the F-statistic lies below the lower-bound critical value, then there is no cointegration among the variables.

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If the F-statistic lies above the upper-bound critical value, then the variables are cointegrated. If the F-statistic lies between the upper- and lower-bound critical values, then the results are inconclusive, and further investigation is needed.

If a long-run relationship exists among the variables, then the ARDL estimates for the lagged level variables in Equation (19) are not spurious. Thus, the long-run coefficients can be calculated. In this calculation, we assume that all the differenced variables in Equation (19) are zero in the long run, i.e., steady state. Thus, the long-run equation corresponding to Equation (19) is as follows:

$$rem_t = \alpha_0 + \alpha_1 y_t + \alpha_2 el_t + \alpha_3 p_t + \alpha_4 expl_t + e_t, \quad (20)$$

where $\alpha_0 = \frac{\beta_0}{-\beta_1}, \alpha_1 = \frac{\beta_2}{-\beta_1}, \alpha_2 = \frac{\beta_3}{-\beta_1}, \alpha_3 = \frac{\beta_4}{-\beta_1}, \alpha_4 = \frac{\beta_5}{-\beta_1},$.

The short-run equation corresponding to Equation (20) is as follows:

$$\Delta rem_t = \beta_0 + \sum_{i=1}^{p1} \gamma_i \Delta rem_{t-i} + \sum_{i=0}^{p2} \delta_i \Delta y_{t-1}^d + \sum_{i=0}^{p3} \theta_i \Delta l_{t-1}^f + \sum_{i=0}^{p4} \vartheta_i \Delta p_{t-i} + \sum_{i=0}^{0} \pi_i \Delta expl_{t-i} + ECT_{t-1} + \epsilon_t, \quad (21)$$

where $ECT_t = rem_t - (\alpha_0 + \alpha_1 y_t^d + \alpha_2 l_t^f + \alpha_3 p_t + \alpha_4 expl_t + e_t)$.

7. Results and Discussion

7.1. Unit Root Test

The results of the ADF unit root tests are shown in Table 3.

Table 3. ADF test results for the variables.

	ADF unit root test						
	Level				First difference		
Variables	**t-stat**	**C**	**T**	**k**	**t-stat**	**C**	**T**
rem_t	-1.997	x		0	-5.786[a]	x	
y_t^d	-2.813		x	0	-5.351[a]	x	
p_t	-2.723		x	0	-5.802[a]	x	
l_t^f	-3.545[b]		x	2	-3.782[a]	x	
Phillips-Perron (PP) unit root test							
rem_t	-2.001	x			-5.695[a]	x	
y_t^d	-3.156		x		-5.293[a]	x	
p_t	-2.723		x		-6.137[a]	x	
l_t^f	-2.896[c]	x			-3.812[a]	x	

Source: Authors' estimates.

Notes: The maximum lag order is set to three, and the optimal lag order (k) is selected based on the Schwarz criterion in the tests. The superscripts a, b and c indicate rejection of the null hypotheses at the 1%, 5% and 10% significance levels, respectively. The critical values for the tests are taken from MacKinnon (1996). The final unit root test equation can include an intercept (C), an intercept and trend (T) or none. In the table, an x indicates that the corresponding option is selected in the final unit root test equation. The sample spans the period from 1973 to 2019.

The null hypothesis of unit root for the log levels of all variables cannot be rejected at conventional levels of significance as the sample t-statistics are less than the critical value in absolute terms. For el, the null hypothesis cannot be at the 1% significance level in the case of the ADF test and at the 5% significance level in the case of the PP test. Consequently, there is still an indication of the unit root process for el. Additionally, the graphical illustration of the variable in Panel A of Figure 4 suggests more of a unit root process rather than a trend stationary or level stationary process. For the first difference of the variables, the null hypothesis of a unit root is rejected at the 1% significance level, as the sample t-statistics are greater than the critical values in absolute terms. Thus, we conclude that all the variables included in the analysis are non-stationary in their log levels and stationary in their growth rates (i.e., they are I[1]).

7.2. Cointegration Analysis Results

7.2.1. Johansen Cointegration Test Results

To apply the Johansen cointegration test, we first estimate a VAR model and then transform it into a VECM. Our time period is from 1970 to 2019, meaning that we have 50 annual observations. We therefore specify a general unrestricted VAR with three lags, as the time-dependent lag selection rule of $4(T/100)^{2/9}$ suggests 3.4 lags.[4] Our sample for estimations and testing therefore runs from 1973 to 2019.

As we mention in Section 5, *expl* only has values for 2017 to 2019, as the expatriate levy was implemented by the government in 2017. Thus, we cannot include it as an endogenous variable in the VAR because it would be impossible to run the regression. The endogenous variables in our VAR analysis are *rem*, *y*, *p* and *el*. Our set of exogenous variables includes the contemporaneous values and two lagged values of *expl* and two dummy variables.[5] Remittance outflows from Saudi Arabia have substantially declined since 2016, owing to fiscal and energy price reforms and low oil prices (see Figure 2). Thus, we include in the VAR analysis a dummy variable taking a value of unity in 2016 and a value of zero otherwise. This variable captures the effects of these reforms and low oil prices on remittance outflows. The second dummy variable captures the spike in *rem* in 1978, which is mostly caused by high oil prices and revenues.

Table 4. VAR lag order selection criteria.

Lag	LogL	LR	FPE	AIC	SC	HQ
0	-28.29	NA	0.0001	2.23	3.17	2.58
1	191.21	345.60*	1.94e-08*	-6.43*	-4.86*	-5.84*
2	201.99	15.14	2.54e-08	-6.21	-4.01	-5.38
3	216.61	18.03	2.95e-08	-6.15	-3.32	-5.09

Source. Authors' estimates.

Note: * indicates the optimal lag order. LogL = log likelihood; LR = likelihood ratio; FPE = final prediction error; AIC = Akaike information criterion; SC = Schwarz information criterion; HQ = Hannan Quinn information criterion.

Table 4 reports the results of the lag order selection. All the information criteria collectively agree that one lag order is preferable. We estimate the VAR with one lag, and it is well-behaved in terms of the stability test and residual diagnostics tests. The residual diagnostic tests include the serial correlation LM test, normality tests and the heteroskedasticity test for the residuals. The results of these tests are reported in Table 5.

Table 5. VAR residual diagnostics and cointegration test results.

Panel A: Serial correlation LM test[a]

Lags	LM-statistic	P-value
1	19.38	0.25
2	10.35	0.85

Panel B: Normality test[b]

Statistic	χ^2	d.f.	P-value
Skewness	4.95	4	0.29
Kurtosis	7.63	4	0.11
Jarque-Bera	12.58	8	0.13

Panel C: Heteroskedasticity test[c]

White	χ^2	d.f.	P-value
Statistic	151.32	130	0.10

Panel E: Johansen cointegration test summary

Data trend:	None	None	Linear	Linear	Quadratic
Test type:	(a) No C and t	(b) Only C	(c) Only C	(d) C and t	(e) C and t
Trace:	1	1	1	2	2
Max-eig:	1	1	1	2	2

Panel F: Johansen cointegration test results for type (b)

Null hypothesis:	$r = 0$	$r \leq 1$	$r \leq 2$
λ_{trace}	109.58***	30.81	12.49
λ^a_{trace}	100.25***	28.19	11.43
λ_{max}	78.77***	18.32	9.36
λ^a_{max}	72.06***	16.76	8.56

Panel D: VAR stability test

Roots	Modulus
0.96	0.96
0.90 - 0.14i	0.91
0.90 + 0.14i	0.91
0.34	0.34

No root lies outside the unit circle. VAR satisfies the stability condition

Source: Authors' estimates.

Notes: [a] The null hypothesis in the serial correlation Lagrange multiplier (LM) test is that there is no serial correlation at lag order h of the residuals. [b] We perform a system normality test with the null hypothesis that the residuals are multivariate normal. [c] The White heteroskedasticity test takes the null hypothesis of no cross-term heteroskedasticity in the residuals. χ^2 is Chi-squared, and d.f. stands for degrees of freedom. C and t indicate the intercept and the trend, respectively. r is the rank of the matrix Π, that is, the number of cointegrated equations. λ_{trace} and λ_{max} are the trace and max-eigenvalue statistics, respectively, and λ^a trace and λ^a max are their adjusted versions. *** denotes rejection of the null hypothesis at the 1% significance levels. The critical values for the cointegration test are taken from MacKinnon, Haug, and Michelis (1999).

Since all the diagnostic tests are satisfied for the VAR, we transform it into a VECM. We employ the Johansen cointegration test to check whether the variables under investigation are cointegrated. We exclude the exogenous variables mentioned above, and we use the critical values from MacKinnon, Haug, and Michelis (1999) in the cointegration test as they are more accurate than previously suggested critical values. The results are reported in Panel E of Table 5. The trace and maximum eigenvalue statistics both reject the null hypotheses of no cointegration. The first three test equations suggest that there is no more than one cointegrating relation. When linear and quadratic trends are included in the test equations (i.e., in the cointegrating space), the results suggest two cointegrated relationships.

The time trend variable should be carefully incorporated as an explanatory variable in the regression analysis with other non-stationary variables. Econometric issues can arise if this variable is not treated appropriately (Nelson and Kang 1984; Mankiw and Shapiro 1985, 1986; Durlauf and Phillips 1988; Pesaran, Pierse, and Kumar

1989). It is preferable to avoid including a linear trend in the long-run part of the VECM. Exceptions are if it produces statistically and theoretically coherent results or if there is a specific reason for its inclusion. In this regard, we note that including a time trend in the long-run part of the VECM changes the signs of the explanatory variables incorrectly. These variables are also no longer statistically significant. Hence, we ignore the last two test options.

Additionally, the test type (a), in which the intercept is omitted from the long-run relationship is not generally considered, as many economic relationships have initial values. One should have a specific reason for considering this option, such as modeling the Fisher equation or an absolute purchasing power parity relationship. Otherwise, this option is generally dropped.

We therefore drop the first and last two options and consider options (b) and (c) in Panel E of Table 5. These options both indicate only one cointegrating relationship among the variables. Further investigation suggests that option (b) is preferable to option (c), as the former produces more reasonable theoretical and statistical results than the latter. Panel F of Table 5 reports the sample values of the trace and maximum eigenvalues from the cointegration test type of (b). For further robustness, we apply the small sample bias correction developed by Reinsel and Anh (1992) and Reimers (1992) to the trace and maximum eigenvalues. Even after applying the small sample bias correction, the test statistics suggest that the variables have no more than one cointegrating relationship.

After concluding cointegration among the four endogenous variables in the VAR, we replace *expl* from the exogenous part of the VECM to the endogenous part of it, given that all the information criteria in Table 4 suggest not more than one lag order would be optimal and a VAR with one lag can accommodate *expl* as an endogenous variable. The purpose is to test whether *expl* establishes cointegration with the other variables. We again test for cointegration for all the possible options and find that, according to the trace statistics, there is only one cointegrating relationship among these five variables in all the five test types. The maximum eigenvalues statistics suggest one cointegrating relationship for options (a), (b), and (c) and two cointegrating relationships for options (d) and (e). We then further test for option (b), the preferred option, using the small sample bias correction as we did above. In the case of the five variables (i.e., *rem*, *y*, *el*, *p* and *expl*), the trace sample values are 127.55, 46.76, 22.95 and the maximum eigenvalues values are 80.79, 23.81, 11.84 for the null hypotheses of $r=0$, $r \leq 1$, $r \leq 2$, respectively. The corresponding critical values for the trace and maximum eigenvalues statistics are 76.97, 54.08, 35.20 and 34.81, 28.59, 22.30, respectively for the 5% significance level. In the small sample bias correction, we consider that there are five endogenous variables in the VAR being estimated with an optimal lag order of one and 47 observations. This makes the adjustment/correction coefficient (47-5*1)/47. Applying this coefficient to the sample values yields 113.98***, 41.79, 20.51 for the trace statistic and 72.20***, 21.28, 10.58 for the maximum eigenvalue statistics. Both the standard and adjusted values of the trace and maximum eigenvalues still indicate only one cointegrating relationship among the five variables, not only at the 5% significance level, but also at the 10% and 1% significance levels. This result allows us to consider single-equation-based or residual-based cointegration methods and long-run estimators. For example, we can use the bounds testing approach based on the ARDL estimation suggested by Pesaran, Shin, and Smith (2001). This method is proven to be superior to other long-run estimators for small sample sizes.

7.2.2. ARDL Cointegration Test Results

As in the VAR analysis, we include *rem*, *y*, *el* and *p* in the dynamic part of the ARDL equation. We use *expl*, *DP2016* and *DP1978* as fixed regressors. Again, as in the VAR analysis, our unrestricted ARDL equation includes three lags of the endogenous variables and two lags of *expl*. We also include their contemporaneous values. Then, following Pesaran and Shin (1999) and Pesaran, Shin, and Smith (2001), we select the optimal lag order for the variables based on the Schwarz information criterion. ARDL(2,0,0,1) is selected, and only contemporaneous values of *expl* are statistically significant alongside the dummy variables. The final ARDL specification successfully passes post-estimation tests of serial correlation, normality, heteroskedasticity, the autoregressive conditional heteroskedasticity (ARCH) effect and functional misspecification. The results of these tests are shown in Table 6.

Table 6. Residual diagnostic test results.

Test	F-statistic	P-value
Normality Test	2.684*	0.261
Breusch-Godfrey Serial Correlation LM Test:	0.906	0.413
Heteroskedasticity Test: ARCH	0.116	0.736
Heteroskedasticity Test: White	0.563	0.818
Ramsey RESET Test	0.206	0.653

Source: Authors' estimates.

Note: * indicates Jarque-Bera test statistic not the F-statistic.

Since the final ARDL specification passes the diagnostic tests, we use it to test for cointegration. We include the intercept in the cointegration test. The results are reported in Panel A of Table 7.

Panel A indicates that the sample F-statistic (11.624) is greater than the upper-bound critical value of Pesaran, Shin, and Smith (2001) for the 1% significance level (4.08). The sample F-statistic is also greater than the upper-bound critical value of Narayan (2005) for the 1% significance level. This result holds regardless of whether we use 50 or 45 observations (5.328 and 5.412, respectively) and addresses the small sample bias. In other words, the null hypothesis of no cointegration can be rejected at the given significance levels. This result suggests that the variables are cointegrated.

We also perform the following robustness check. As reported in Panel A, the Schwarz information criterion does not select more than two lags in the final ARDL specification search. Thus, we also estimate an unrestricted ARDL equation with two lags as the maximum lag order for each variable. However, in this equation, we include *expl* as a dynamic regressor. The final ARDL(2,0,0,1,0) specification is selected based on the Schwarz information criterion (see Panel B). The Schwarz information criterion selects the same lag orders for the variables, including *expl*, as were previously selected. However, we do not treat the latter as a fixed regressor in this equation. The post-estimation test results indicate that the final specification of ARDL(2,0,0,1,0) has no issues regarding serial correlation, normality, heteroskedasticity, the ARCH effect and misspecification. These results are available from the authors on request. We perform the bounds test again, and the results are reported in Panel B of Table 7. The null hypothesis of no cointegration can be rejected at any conventional significance level. This result holds even if we use upper-bound critical values from Narayan (2005). Thus, based on the two ARDL analyses, we conclude that the variables under consideration are cointegrated.

Table 7. Results of the bounds test for cointegration.

Panel A. **Specification used: ARDL(2,0,0,1)** **Null hypothesis of no cointegration** $H_0: \beta_0=\beta_1=\beta_2=\beta_3=\beta_4=0$				**Panel B.** **Specification used: ARDL(2,0,0,1,0)** **Null hypothesis of no cointegration** $H_0: \beta_0=\beta_1=\beta_2=\beta_3=\beta_4=\beta_5=0$		
Sample F-statistic	11.604			Sample F-statistic	9.828	
Sample size	47			Sample size	47	
Critical value						
Lower bound	Upper bound	SL		Lower bound	Upper bound	SL
Asymptotic: n=1,000				Asymptotic: n=1,000		
2.37	3.20	10%		2.20	3.09	10%
2.79	3.67	5%		2.56	3.49	5%
3.15	4.08	1%		3.29	4.37	1%
Finite sample: n=50				Finite sample: n=50		
2.538	3.398	10%		2.372	3.320	10%
3.048	4.002	5%		2.823	3.872	5%
4.188	5.328	1%		3.845	5.150	1%
Finite sample: n=45				Finite sample: n=45		
2.560	3.428	10%		2.402	3.345	10%
3.078	4.022	5%		2.850	3.905	5%
4.270	5.412	1%		3.892	5.173	1%

Source: Authors' estimates. Note: SL = significance level; n = number of observations.

7.3. Long-run Estimates

Table 8 reports the results of estimating the long-run ARDL(2, 0, 0, 1, 0) specification.

As the table shows, all the explanatory variables are statistically significant at conventional levels. Their signs are coherent with the theoretical discussion in Section 2. We discuss these findings in the next section.

Table 8. Long-run estimates.

Variable	Coefficient	Std. error	t-statistic	Prob.
y_t^d	1.697	0.424	3.997	0.000
el_t^f	0.998	0.537	1.858	0.071
p_t	-1.426	0.486	-2.935	0.005
$expl_t$	-0.092	0.037	-2.452	0.019
c	-14.799	5.606	-2.640	0.012

Source: Authors' estimates.

7.4. Short-run Estimates

For the short-run estimation, we perform ECM modeling in the framework of the general-to-specific methodology mentioned in Section 6. The general unrestricted ECM specification includes an intercept, two lags of the dependent and independent variables and contemporaneous values of the independent variables, and ECT with one lag. The changes in the dummy variables $DP1978$ and $DP2016$ are included in the general unrestricted ECM, following the standard methodology for the transformation of variables, including deterministic regressors, from a level equation to a growth rate equation (e.g., Juselius [2006]). We also include the dummy variable $DP1986$ because of a large drop in the residuals of the general unrestricted ECM in 1986. The oil price fell substantially in that year, resulting in decreased government revenues and an economic recession. This drop causes a misspecification issue, as the Ramsey RESET test shows. The dummy variable takes a value of unity in 1986 and a value of zero otherwise. ECT is calculated as follows:

$$ECT_t = rem_t - (1.697 * y_t^d + 0.998 * l_t^f - 1.427 * p_t - 0.092 * expl_t - 14.799).$$

Once we form the general ECM specification, we exclude insignificant variables while performing the residual diagnostic tests and misspecification tests. Doing so ensures that the resulting ECM specification is well-behaved. The final ECM specification is presented in Panel A of Table 9, and diagnostic test results are reported in Panel B of Table 9.

The final specification successfully passes all the residual diagnostic tests. These tests include the Jarque-Bera statistic for the normality of the residuals and the LM test for serial correlation. We also perform the White and ARCH tests for heteroskedasticity and the Ramsey RESET test for specification error. In addition, the estimated coefficients are statistically significant at conventional levels, and their signs match theoretical expectations. We discuss the short-run findings in the next section.

7. Results and Discussion

Table 9. Final ECM estimation results.

Panel A: Short-run coefficient estimates

Variable	Coefficient	Std. error	t-statistic	P-value
ECT_{t-1}	-0.204	0.040	-5.132	0.000
C	0.015	0.017	0.864	0.393
Δrem_{t-1}	0.198	0.060	3.322	0.002
Δy_t^d	0.764	0.165	4.640	0.000
Δl_{t-1}^f	0.264	0.156	1.689	0.100
Δp_t	-1.337	0.094	-14.158	0.000
Δexp_t	-0.046	0.015	-3.049	0.004
$\Delta DP1978$	0.212	0.063	3.376	0.002
$\Delta DP2016$	-0.214	0.084	-2.541	0.015
$DP1986$	-0.309	0.097	-3.185	0.003

Panel B: Diagnostic test results

Jarque-Bera normality test	1.001	0.606
	F-statistic	P-value
Breusch-Godfrey serial correlation LM test:	0.504	0.608
Heteroskedasticity test: ARCH	0.601	0.442
Heteroskedasticity test: White	0.388	0.934
Ramsey RESET Test	0.282	0.598

Source: Authors' estimates.

Notes: The dependent variable is Δrem_t.

8. Discussion of the Empirical Findings

The unit root test results in Table 3 show that the remittance outflow, GDP, foreign employment, and the price level are non-stationary.[6] In other words, and their mean, variance and covariance change over time. Non-stationarity means that any shocks to these variables, including policy-related shocks, may cause permanent changes. Non-stationary variables that drift over time may share a common stochastic trend. The cointegration test results in Tables 5 and 7 indicate that the variables in this study share such a trend. Thus, the remittance outflow, GDP, foreign workers, the price level and the expatriate levy establish a long-run relationship that can be interpreted using economic theory. Estimating numerical values for this relationship may help understand the long-run behavior of the remittance outflows from Saudi Arabia. Decision-makers may consider this information when taking relevant policy measures.

First, all four variables have statistically significant and the theoretically expected impacts on the remittance outflow, as discussed in Section 4. More precisely, the estimated elasticities of remittance outflows with respect to GDP and non-Saudi employment are positive. A 1% increase in Saudi Arabia's GDP will increase remittance outflows by 1.7% in the long run, keeping other factors unchanged. Similarly, a 1% increase in non-Saudi employment will increase remittance outflows by 1.0%. The positive impact of non-Saudi employment on remittances is intuitively straightforward. When the number of migrant employees increases, their remittances should increase as well. The opposite relationship could theoretically arise if an increase in the number of migrant workers creates strong competition in the Saudi labor market. In that case, the wage level could decline so dramatically that remittance outflows would decrease. This situation does not hold in Saudi Arabia, however. According to SAMA (2020), the average wage increased from 1,359 riyals in 2005 to 3,041 riyals in 2019. The CPI-deflated average real wage at 2010 prices increased 1.6 times from 2005 to 2019.

Several other explanations for the positive and significant relation between the volume of foreign workers and remittance outflows exist. They include financial support for family members left behind and a lack of investment opportunities in Saudi Arabia. Migrant workers also lack ownership and property rights in Saudi Arabia and have no opportunities for permanent residence. Finally, the return on investment in their home countries may be high. According to Lucas and Stark (1985), altruism, that is, financial support for family members left behind, is the most straightforward of these reasons. They also cite the desire to inherit, the preference for investing in assets in one's own country and the intention to return home. These desires also motivate migrant workers to send money to their countries of origin.

We find that an increase in the income level is associated with an increase in remittance outflows. This result suggests that macroeconomic conditions in the Kingdom are a crucial driver of the outflow of remittances. We have two main explanations for this finding. First, an increase in the income level expands economic activity, which, in turn, increases the demand for labor, including non-Saudi labor. Statistically speaking, Saudi Arabia's GDP grew more than five times during the study period from 1970 to 2019. Non-Saudi employment increased by more than 22 times during that time. Increased employment of migrant workers leads to a greater remittance outflow. Second, an increased income level increases the average wage level, as we discuss above, and this wage growth, in turn, increases remittance outflows.

8. Discussion of the Empirical Findings

Table 8 shows that the price level and the expatriate levy negatively affect remittance outflows from Saudi Arabia. Numerically, a 1% increase in the Kingdom's price level leads to a 1.4% decrease in remittance outflows. By contrast, a 1% increase in the expatriate levy reduces remittance outflows by 0.1%. We expect the price level, as a measure of the cost of living, to exert a negative impact on remittance outflows. Migrant workers cannot send more money back to their home countries if the cost of living increases in the Kingdom. The price level we use in this study also measures production costs. An increase in the production cost also implies a negative effect on remittance outflows. Such an increase forces producers to either lower the number of workers or reduce their compensation to maintain profitability. In such a circumstance, producers might opt to reduce the number of foreign workers they employ because doing so also lessens the burden of the expatriate levy that they must pay for their foreign workers. As Figure 2 shows, the aggregate price level, as measured by the GDP deflator, has increased considerably in Saudi Arabia. Numerically, it grew about 23 times from 1970 to 2019. Its negative association with remittance outflows is therefore notable.

We find that the expatriate levy negatively impacts the remittance outflow, although the magnitude of the estimated coefficient is small. The negative relationship between this fee and remittance outflows does not require an explanation, as the fee acts as a tax on foreign workers. The greater the fee is, the less money foreign workers receive to send back home. The small magnitude of the elasticity may arise for several reasons. First, our analysis covers the period from 1970 to 2019, and the expatriate levy was implemented by the Saudi government in 2017 (FBP 2017). In other words, the variable has only three values for the last three years of the 50-year sample, and, thus, the impact of this variable on remittance outflows has little variability. Second, some public and private entities pay the whole amount of the expatriate levy, whereas others pay a certain percentage of the levy on behalf of their foreign workers in the Kingdom. These payments diminish the negative impact of the levy on foreign workers' earnings and, thus, on their remittances to their home countries. Third, a levy must be paid for each non-Saudi family member. Most foreign workers in the Kingdom are either single or live with a wife.

We next discuss the empirical findings from the short-run analysis. First, Table 9 shows that the coefficient of ECT with one lag is negative and highly statistically significant. The negative value indicates that a short-run disequilibrium corrects back to the equilibrium level. The high statistical significance level shows that the remittance outflow's long-run relationship with real GDP, non-Saudi employment, the price level and the expatriate levy is stable. Numerically speaking, 20.4% of a short-run disequilibrium is corrected back to the equilibrium level in the subsequent year. The full correction takes 4.9 years. The correction is halfway complete after 3.4 years,[7] and the second half of the full restoration takes 1.5 years. Once the correction process starts, the first half of the correction takes more time than the second half does.

Additionally, the short-run estimations show that GDP, price and the expatriate levy have contemporaneous impacts on the outflow of remittances while foreign workers have a lagged effect. Numerically, a one percentage point increase in the growth rate of Saudi GDP increases the growth rate of remittance outflows by 0.8 percentage points. A one percentage point increase in the growth rate of the price level reduces the remittance outflow's growth rate by 1.3 percentage points. A one percentage point increase in the expatriate levy's growth rate reduces that of remittance outflows by 0.05 percentage points. Finally, a one percentage point increase in the growth rate of foreign employment in the previous (present) year increases the remittance outflow's growth rate by 0.3 percentage points in the present (next) year. The intuition for these results follows the same reasoning as we discussed for the long-run estimation results.

Remittance outflows exhibit a dynamic relationship. A one percentage point increase in the remittance outflow's growth rate in the previous year accelerates its current growth rate by 0.2 percentage points. Lastly, all three dummy variables have economically reasonable impacts on the growth rate of remittance outflows. The high oil price in 1978 has a statistically significantly positive effect on the growth rate of remittance outflows. The low oil price and the resulting economic recession in 1986 have a statistically significantly negative effect. The energy price and fiscal reforms coupled with the low oil price in 2016 similarly have statistically significantly negative impacts. Numerically, keeping other factors constant, the average growth rate of remittance outflows increased by 0.2 in 1978. It decreased by 0.2 and 0.3 in 2016 and 1986, respectively. These are expected findings, as economic activity, the income level and, consequently, remittance outflows from Saudi Arabia are considerably influenced by oil revenues. The effects of these shocks are ultimately reflected in remittance outflows. For example, the highest growth rate of remittance outflows over our study period, 88.8%, was recorded in 1978 (World Bank 2020b). It would be useful to know how much the outflow of remittances changed in percentage terms due to the events that occurred in the given years, which are represented by the dummy variables. For this, we use a formula suggested by Halvorsen and Palmquist (1980), that is, *exp(the coefficient of a dummy variable)*$^{-1}$, and considered small sample bias correction for the estimated coefficient of a dummy variable, that is, *exp(the coefficient of a dummy variable) is replaced with exp(the estimated coefficient of a dummy variable-(-estimated variance of the coefficient/2))*, suggested by Kennedy (1981). Accordingly, we calculate that the outflow of remittances changed by 0.24, -0.19, and -0.27 percentage points due to the events that occurred in 1978, 1986, and 2016, respectively.

9. Conclusion and Policy Insights

The outflow of remittances is a key element of the balance of payments and has implications for the macroeconomic environment of the host country. In terms of the remittance outflow, Saudi Arabia is a leading economy globally and ranks first in the GCC region until 2015. Thus, understanding the long-run and short-run behaviors of remittance outflows is very important. This importance is also justified by the fact that remittance outflows are a key source of leakages from the Saudi economy. These leakages mean that Saudi Arabia's fiscal spending multipliers are smaller than those of similar economies, as previous studies show. We investigated how theoretically predicted determinants shape remittance outflows in the long and short runs. We found that in the long run, domestic economic activity and foreign employment positively affect remittance outflows. By contrast, the domestic price level and the expatriate levy have negative impacts. These determinants also influence remittance outflows in the short run.

This study's findings may be useful for macroeconomic policymaking. As mentioned above, the outflow of remittances is a major channel for leakages of money and income from the country. To overcome these issues, the authorities implemented a nationalization policy in the labor market and imposed a levy on foreign workers and their dependents. These policies aim to encourage businesses to hire Saudis in different economic sectors. We found that the expatriate levy has a negative effect on remittance outflows. In this regard, imposing an expatriate levy seems to be an effective policy as it curbs leakages and might, to some extent, discourage foreign workers from participating in the labor force. However, implementing a higher expatriate levy is not straightforward. If the levy is high enough to discourage foreign workers' participation, a labor shortage may occur. Given that foreign workers comprise around 80% of total employment in the private sector, the labor shortage may increase the wage rate and contract economic activity unless enough Saudi workers are willing to enter the labor market. To this end, policymakers should implement a balanced policy by finding the right tradeoff between the expatriate levy and economic activity. This is a useful topic for future research.

Additionally, a higher levy reduces the disposable income levels of foreign workers, which can reduce their final consumption spending and that of their dependents. It also reduces the profit of the entities hiring them, which can reduce their investment spending. Both outcomes are unfavorable for boosting aggregate demand and, consequently, economic growth.

Another balanced measure that might be considered in policymaking is easing restrictions on owning entities, real estate and investment for non-Saudis and providing them with permanent residency. Studies show that a lack of investment opportunities in the host country is one of the key drivers of outflow remittances. The lack of ownership and property rights for migrant workers in Saudi Arabia is another motivation for them to send remittances to their home countries. Easing these restrictions would encourage foreigners to spend more in the Saudi economy in the form of consumption and investment. Both types of spending will lead to economic growth. In this regard, the recent reforms and initiatives that the government is implementing in line with Saudi Vision 2030 should be acknowledged.

Lastly, when policymakers plan to implement measures to impact the remittance outflow, they may wish to consider the time scale over which these measures will have effects. The outflow of remittances takes five years to fully adjust to its long-run equilibrium relationship with economic activity, prices, foreign employment, and the expatriate levy. Adjusting halfway takes more than three years. In other words, the adjustment period is not short, and the first half of the adjustment process takes longer than the second half does.

Endnotes

1 See, for example, Adams and Page (2005), Pradhan, Upadhyay, and Upadhyaya (2008), Qayyum, Javid, and Arif (2008), Adams and Cuecuecha (2010), Javid, Arif, and Qayyum (2012) and Ratha (2013).

2 One can use the GDP deflator or the consumer price index (CPI) as a measure of the price level. We employ the former for the following reasons. First, the theoretical model does not specify whether the price level is the GDP deflator or the CPI. Thus, both measures can be considered in the empirical analysis. Second, CPI data are available for 1980 in the best-case scenario, but our study starts in 1970. Hence, using CPI would cost us the loss of information from the first 10 years. Third, previous studies on outflow remittances from Saudi Arabia also prefer the GDP deflator to the CPI (e.g., Al-Abri, Genc, and Naufal [2018]). Additionally, the GDP deflator reflects prices faced by households, firms, and the government, while the CPI only reflects prices for households. The outflow remittances are affected not only by consumer/household prices but also by producer prices. Moreover, the Saudi government implemented domestic energy price and fiscal reforms (including the introduction of a value added tax and expatriate levy among other initiatives). These reforms can exert adverse effects on outflow remittances through their impacts on the cost of living and prices/cost of producing goods and services.

3 The expatriate levy ($expl$) is not shown in the graphs because it has only three observations (i.e., for 2017 to 2019).

4 In the formula, T is the number of observations.

5 We include two lagged values of $expl$ because this variable has only three values. Hence, we technically cannot have more than two lags of this variable.

6 Although we cannot run formal unit root tests on the expatriate levy variable, the graphical illustrations of the log level and the growth rate of the variable are in favor of the conclusion that the former is non-stationary while the latter is stationary.

7 This value is calculated as ln(2) divided by the absolute value of the speed of adjustment coefficient (SOA). Alternatively, we can calculate the time required for halfway correction as ln(05)/ln((SOA+1), which yields 3.0 years (e.g., see Kim, Ogaki, and Yang [2003]; Foldvari [2012] for a discussion of the method).

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Notes

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About the Project

This study was part of the KGEMM Research and Policy Studies project. The project aims to leverage the work done in developing the KGEMM model to produce policy and research studies that can help Saudi Arabian decision-makers enhance their understanding of the Kingdom's domestic and international macroeconomic-energy linkages.

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