Macroeconomic effects of grain price volatility in Morocco

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Abstract

This paper investigates how cereal price volatility impacts import bill, tax revenue and foreign exchange reserves in Morocco. It uses GARCH family models to characterize the price and exchange rate volatility functions, ARDL model and Toda and Yamamoto's (1995) causality test to study respectively cointegration and causal relationship. Based on monthly data between January 1999 and December 2019, we find that 1% increase of price volatility and volatility-import leads to respectively increase the import bill by 0.07% and 16.7% on the long run. Meantime, the short-run estimates suggest that the effects of price volatility and the volatility-import level are negative meaning that the lagged value of these variables will have a positive impact on the next month's import bill. Thus, we assume that price volatility should be heavier on the import bill when the annual production is low. Our results also indicate that cereals price volatility can induce serious consequences because it directly causes an increase in the overall import bill and indirectly influences import tax revenues and foreign exchange reserves, especially when it is associated with a poor domestic harvest.

Keywords: *Price volatility, Grain imports, Morocco.*

1. Introduction

The cereals sector is one of the most strategic sectors in the world economy, especially thanks to its crucial role for food security, a top political priority for all countries. The current context of Ukrainian-Russian conflict and the Covid-19 crisis have induced strong pressure on global supply chains that led several countries to increasing grain imports and setting up a cereal strategic stock.

In Morocco, the cereals sector stands for the backbone of the country's agri-food system with a contribution of 1 to 3% to the GDP and 10 to 20% to the Agricultural GDP (MAFRD-WF, 2018). However, this production does not guarantee and preserve self-sufficiency of the country. The latest report published by OECD and FAO (2020) designates Morocco as one of the world's major cereal importing countries. The imported quantities vary between 3.8 and 7.5 million tons annually. They are dominated by soft wheat (27 to 55% of imports), followed by maize (27% - 43%), durum wheat (9% - 20%) and barley (2% - 15%) (Figure 1). For example, the decrease registered in 2006-2007

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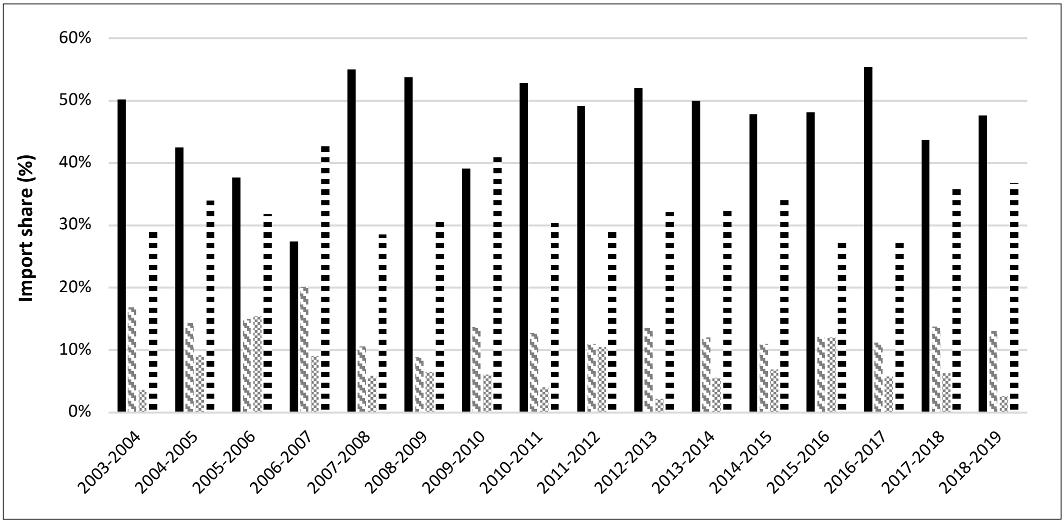
Figure 1 - Evolution of the Moroccan cereal imports share by cereal product.

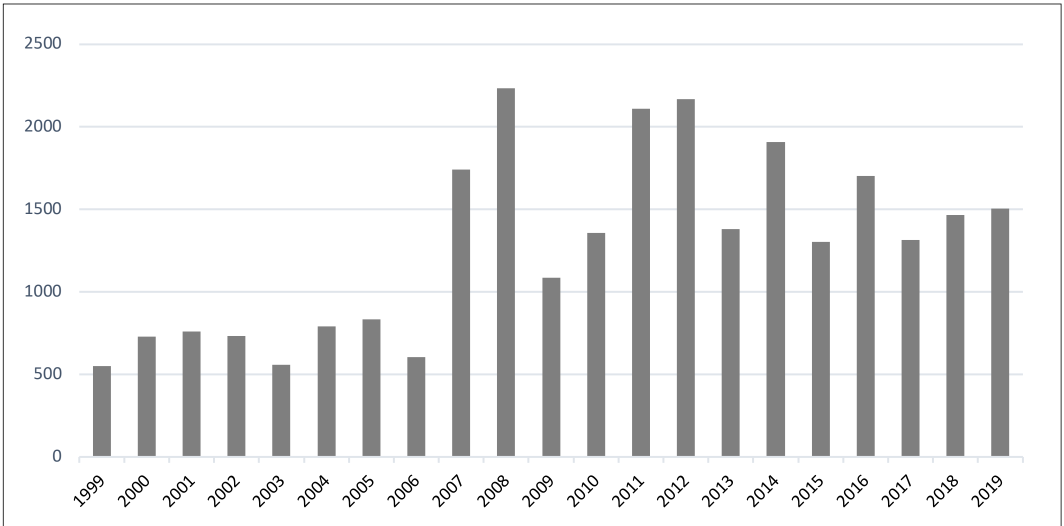


Source: NIOCL, 2020.

and 2017-2018, especially for soft wheat, was the result of a good crop year (NIOCL, 2020).

On the other hand, the Moroccan cereal imports bill has increased by 173% over the past 20 years, according to the Exchange Office of Morocco (2020), going up from nearly USD 550 million in 1999 to nearly USD 1,505 million in 2019 (Figure 2). Many popular factors explain this variation such as the variability of the local production (especially from drought and climate variability), the country's supplying grain policy and the international market prices (Jouamaa, 2021). Indeed, the increase in the economic bill registered in 2007-2008 and 2011-2012 have occurred as a result of commodity price spike in the international market. This situation has become significantly higher and more volatile over the past decade (especially during the Covid-19 pandemic and the Ukrainian crisis).

Figure 2 - Moroccan cereal imports bill (million USD).



Source: Exchange Office of Morocco (2020) and International Financial Statistics data.

Considerable attention has been paid in empirical economics to testing for the effects of commodity price volatility by studying economic growth, interactions with financial markets or on other macroeconomic indicators, such as external debt, inflation, interest rate or public finance. However, very little research has so far focused on the volatility of grain prices in Morocco by taking import tax revenue, foreign exchange and the import bill as variables of interest. This paper aims at contributing to fill this gap by providing relevant information on macroeconomic effects of grain price volatility based on econometric analysis. To fulfill this objective, the first section of this paper gives a brief summary of the relevant literature. Section 2 presents the methodological approach that is suited to answer raised questions. Research results and concluding remarks are reported in subsequent sections.

2. Literature review

Economists have long debated the effects of commodity price volatility on several macroeconomic indicators. The existing studies generally have focused either on the incidence of a price on macroeconomic variables (trade balance, economic growth, interest rate or money supply) or on modeling commodity price fluctuations. Deaton (1999) calculated the price index of 25 commodities and found that there is a close positive relationship between commodity price volatility and growth in Sub-Saharan Africa. Ocran and Biekpe (2007) examined, through autoregressive vector model (VAR) and Granger causality, the effects of commodity prices shaped by monetary policy in South Africa. They concluded that there is some merit in using commodity prices (particularly the average gold price and the IMF's metals price index) as informational variables in setting monetary policy. Hegerty (2016) used a VAR(1)-Multivariate GARCH model to identify spillovers among important commodity prices and output, inflation, exchange and interest rates in major emerging markets. While each commodity and each country behaves differently, Hegerty (2016) found that Chile is most closely tied to the copper price, and Indonesia to oil and tin prices, while other countries such as Brazil, Russia and the Philippines are less affected.

Besides, Céspedes and Velasco (2012) showed that commodity price shocks have a significant impact on output and investment dynamics. This impact tends to be larger for economies with less developed financial markets. The analysis leads also to the finding that international reserve accumulation, more stable political systems, and less open capital accounts tend to reduce the real exchange rate volatility in episodes of large commodity price shocks. On their part, Ehrhart and Guerineau (2013) carried out investigations on 80 developing countries over 1980-2008 period with the aim of studying the impact of commodity price volatility on tax revenues. They measure the price volatility of 41 commodities in the sectors of agriculture, minerals and energy through computing a country-specific index. Other explanatory variables affecting tax revenue such as the GDP per capita, structure of the economy, degree of openness, exchange rate and tax administration capacity are included as well. The results show that tax revenues in developing countries are hurt by the volatility of commodity prices. Indeed, the volatility of export prices reduces income taxes while the volatility of import prices decreases revenues from international trade taxes. Using a multiple-equation panel data model, Boere *et al.* (2015) examined the effects of farm gate price volatility on agricultural land-use changes between 2000 and 2013. Their results showed that price volatility generates ambiguity about the difference between actual and expected output prices and leads to a change in the producer's decision. El-Karimi and El Ghini (2020) used a Structural Vector Autoregression (SVAR) model to examine the transmission effects from world food commodities prices to Morocco's food inflation between January 2004 and December 2018. They argued that global food price shocks have a significant positive impact on the country's food inflation. They found that the transmission effect varies across commodities. A positive external food price shock has a significant positive impact on the consumer prices of cereals and oils, while those of dairy and beverages are insignificant.

Other researchers have analyzed the influence of oil price volatility. Ito (2012) examined the effect of oil price changes on inflation, real effective exchange rate and real GDP for Russia using the VAR model and cointegration analysis. The results showed that, in the long run, a 1% increase (decrease) in oil prices contributes to the growth (decline) in real GDP by 0.44%. In the short run, he concluded that rising oil prices not only stimulate inflation and economic growth negatively and positively, respectively, but also induce real effective exchange rate appreciation. Baumeister and Peersman (2013) examined, through a BVAR model, the relationship between oil prices and United States economic performance from 1974 onward. They found that oil supply shocks account for a smaller fraction of real oil prices volatility in more recent periods, unlike oil demand shocks. Baumeister and Kilian (2014) assessed the relationship between oil and food prices. They found that increases in agricultural commodity prices (corn, soybeans, wheat and rice) have contributed little to increases in oil prices, and US retail food price for the small cost share of agricultural products in higher retail food price. Brini *et al.* (2016) examined the impact of oil price shocks on inflation and the real exchange rate in six oil importers and exporters MENA countries: Tunisia, Morocco, Algeria, Bahrain, Saudi Arabia and Iran using a SVAR model. The results showed that the impact of oil price shocks is positive and significant on real exchange rate of the oil-importing countries (Tunisia and Morocco) and insignificant on the inflation variation.

On the other hand, several empirical studies have been carried out to determine the effects of Terms of Trade[1] (ToT) volatility. Cavalcanti *et al*. (2015) studied the impact of commodity terms of trade volatility on economic growth, total factor productivity, physical capital accumulation and human capital acquisition. Using the standard system generalized methods of moments (GMM), the main finding was that commodity ToT volatility exerts a negative impact on economic growth. The results also indicate that the negative growth effects of commodity ToT volatility offset the positive impact of commodity booms. Also, Andrews and Rees (2009) explored the effects of ToT volatility on growth for a sample of 71 countries between 1971 and 2005. They conclude that ToT volatility has positive and significant impact on the volatility of the output growth and inflation. They suggested that the political and structural framework of the markets seem to be the key of the magnitude of these effects. Specifically, the adoption of floating exchange rates and the development of financial markets are associated to lower macroeconomic volatility for economies that are subject to sizeable ToT shocks.

Other authors have examined the consequences of exchange rate volatility. Mckenzie (1998) analyzed the effect of exchange rate volatility on Australian trade flows using ARCH models. The results were inconclusive, and suggested that the impact of exchange rate volatility differs between traded good markets although it remains difficult to firmly establish the nature of the relationship. On their side, Lotfalipour and Bazargan (2014) investigated the impact of real exchange rate volatility on Iran's trade balance using a GARCH model. They found that the exchange rate and the export variable had no significant effect on the trade balance, while the import variable had a negative and significant effect. Vergil (2002) studied the impact of real exchange rate volatility on Turkey's export flows to the United States and its three main trading partners in the European Union between 1990 and 2000. Their results underlined negative effects of the real exchange rate volatility on real exports. However, Ishimwe and Ngalawa (2015) investigated the impact of exchange rate volatility on South Africa's manufacturing exports to the United States between 1990 to 2014. They showed that exchange rate volatility increase has significant positive effects on manufacturing exports on the long run and insignificant effects on the short run.

On the other hand, volatility modeling has been a subject of second research component in financial markets. Batten *et al*. (2010) modeled the monthly price volatility of four precious met-

[1] ToT is the ratio of export to import prices.

als (gold, silver, platinum and palladium) and studied the macroeconomic determinants. They found that the volatility of gold can be explained by currency variables, but not for silver. They also showed that precious metals are too distinct to be considered as a single asset class or represented by a single index. Dönmez and Magrini (2013) examined the main determinants of price volatility for agricultural commodity (wheat, corn and soybean) using the GARCH-MIDAS model between 1986 and 2012. They suggested that modeling agricultural price volatility as a product of high and low frequency components is more efficient than filtering it through a standard GARCH (1,1) model. They found also that supply and demand and speculation indicators are crucial in explaining the low-frequency component of volatility while monetary factors and energy markets play a less important role. In addition, monetary factors (interest and exchange rates) and energy markets (oil price volatility) appear to play a significant but less important role in all agricultural commodity markets. However, the excess of non-commercial activity on the derivatives markets seems to contribute to increasing the low frequency component of wheat and corn volatilities. In a similar research, Kalkuhl and von Braun (2016) found that price spikes are negatively correlated with supply shocks and positively correlated with economic growth shocks. Specifically, energy prices, monetary policy, speculative activities, sudden trade restrictions and lack of information are the important determinants of price volatility in agricultural markets. Cinar (2018) used the BEKK-MGARCH[2] model to test the interdependence between volatility movements and transmissions in grain prices between January 2003 and August 2017. The results confirmed one-way price volatility transmission from corn and barley to the wheat market. Similar method was used by Kakhki *et al.*, (2019) to determine the price volatility transmission between the International barley market, Iran domestic market and Iran Mercantile Exchange. They found a positive interaction between shocks of the world market and Iran Mercantile Exchange. They also indicated that barley price in domestic market is more volatile than other markets. More recently, Ihle *et al.* (2019) investigated the relationship between political instability and the volatility of Palestinian food prices (bananas, milk, onions and pears) between 2004 and 2011. Using a GARCH model, they showed that the varying conflict intensity has modest impact on weekly average prices. Higher return volatility of pears and onions is only partly impacted.

Though various researchers analyzed the effects of commodity price volatility, they focus mostly on one variable or on modeling commodity price fluctuations. This paper is devoted to analyze the relationship between grain price volatility and three Moroccan macroeconomic variables – import bill, tax revenue and foreign exchange reserves. As follows, we present the methodological approach adopted and subsequently analyze the obtained results.

3. Methodology

3.1. *Conceptual framework*

The econometric methods treated in the literature to characterize the effects of volatility on macroeconomic variables fall into two groups. The first group focuses on modeling volatility functions of specific commodities using GARCH family models, which accommodate time varying volatility and the autocorrelation that is frequently evident in financial time series (Batten *et al*., 2010; Dönmez and Magrini, 2013; Cinar, 2018; Ihle *et al.*, 2019). The second group stands for the analysis of price shocks and causal links between volatility variables. The common methods that are used in financial and economic research to examine volatility and macroeconomic spillovers are notably "*Autoregressive Distributed Lag (ARDL)*", "*Panel data*" and "*Vector Autoregressive*" family models (VAR, BVAR, SVAR). On the other hand, the

[2] BEKK-GARCH is the Baba-Engle-Kraft-Kroner (BEKK) version of the multivariate Generalized Autoregressive Heteroskedastic (MGARCH) model.

directions of causation between variables are examined using Granger or Toda and Yamamoto (1995) causality tests.

As mentioned in the literature review, linkages between macroeconomics and commodity price volatility have been documented (Deaton, 1999; Ocran and Biekpe, 2007; Hegerty, 2016). This paper aims to analyze the impact of grain price volatility on import bill, tax revenue and foreign exchange reserves in Morocco using GARCH family methods and the most robust technique of ARDL modeling approach to find the short run and long run estimates of the model.

3.2. *Econometric analysis*

Based on the works of Adam *et al.* (2001), Ehrhart and Guerineau (2013), Keen and Lockwood (2010) and Khattry and Rao (2002), the model is formulated as follows:

$$T_i = \alpha + \beta_1 vol_{imp} + \beta_2 vol_{exp} + \beta_3 I_{imp} + \\ + \beta_4 I_{exp} + \beta_5 V_i + \varepsilon_i \quad (1)$$

Where tax revenue (T_i) depends on commodity prices volatility[3] for imports (vol_{imp}) and for exports (vol_{exp}), commodity import and export price indices (I_{imp}, I_{exp}) and a vector (V) which captures other explanatory variables affecting tax revenues, in particular GDP, the degree of the openness,[4] the share of agriculture in GDP and the exchange rate volatility.

In addition, we include two other variables namely the import bill as an independent variable and tax revenue and foreign exchange reserves as explanatory variables. Our model can be written as follows:

$$Imp_t = \alpha + \beta_1 Imp_{t-1} + \beta_2 Fisc_t + \beta_3 Res_t + \\ + \beta_4 Vol\_gr_t + \beta_5 vol\_ex_t + \beta_6 Int_t + \varepsilon_i \quad (2)$$

Where Imp, Fisc, Res, Vol_gr and Vol_ex respectively stand for imports bill, import tax revenues, foreign exchange reserves, price volatility and exchange rate volatility. We have added a new variable (Int) to the model, in order to take into account the variability of import volumes, which depend mainly on the level of local production. This variable represents the interaction between price volatility and the proportion of cereal imports.

To determine the price and exchange rate volatility function and consider the volatility of each financial series, we first construct our own price index of the main cereals imported by Morocco.[5] Then, we estimate the price volatility function by choosing the most appropriate specification from GARCH (General Autoregressive Conditional Heteroscedastic) models. We follow the Box-Jenkins method to determine the adequate representation of the ARMA (p, q) and apply the formal ARCH test to confirm the absence of autocorrelation of the residuals and the absence of homoscedasticity.[6]

According to Bollerslev (1986), the variance equation of GARCH model can be defined as follows:

$$h_t = \alpha_0 + \sum_{i=1}^{p} \alpha_i \varepsilon_{t-i}^2 + \sum_{j=1}^{q} \beta_j h_{t-j} \quad (3)$$

Where $p \geq 0$ et > 0; $\alpha_0 > 0$; α_i *and* $\beta_i \geq 0, i$;

The Augmented Dickey-Fuller (ADF) and the Philips-Perron (PP) tests are used for each variable to test for the presence of unit roots.

In the next step, we use the cointegration methods and ARDL model (*Autoregressive Distributed Lag*) developed by Pesaran *et al.* (2001) to estimate the coefficients of the long-run cointegrating relationship and the corresponding ECM (Error Correction Model). This approach has several advantages. First, it can be applied irrespective of whether the regressors are I(0), I(1) or mutually cointegrated. Second, it is more robust and efficient with small samples than other cointegration techniques.[7]

[3] The authors measure here commodity price volatility by calculating the standard deviation of the first difference of the price indices for each year based on the twelve monthly price indices.

[4] The degree of openness is measured as the sum of imports and exports as share of GDP.

[5] Our steps of calculating the price index of cereals are explained in Annex 1.

[6] The same method was followed to determine the exchange rate volatility.

[7] It captures the temporal dynamics in the evolution of a time series and corrects the problem of serial correlation and endogeneity (Atozou and Akakpo, 2017; Pesaran et al., 2001).

The equation (2) can be expressed as follows:[8]

$$\Delta imp_t = \alpha_0 + \sum_{i=1}^{n} \alpha_{1i}\, \Delta imp_{t-i} + \\ + \sum_{i=1}^{n} \alpha_{2i}\, \Delta Fisc_{t-i} + \sum_{i=1}^{n} \alpha_{3i}\, \Delta Res_{t-i} + \\ + \sum_{i=1}^{n} \alpha_{4i}\, \Delta Vol\_gr_{t-i} + \\ + \sum_{i=1}^{n} \alpha_{5i}\, \Delta Vol\_ex_{t-i} + \sum_{i=1}^{n} \alpha_{6i}\, \Delta Int_{t-i} + \\ + \beta_1 imp_{t-1} + + \beta_2 Fis_{t-1} + \beta_3 Res_{t-1} + \\ + \beta_4 Vol\_gr_{t-1} + \beta_5 Vol\_ex_{t-1} + \beta_6 Int_{t-1} + \varepsilon_t \quad (4)$$

Where
Δ: the first difference of the variables
α_0: Constant
α_{1i},…, α_{6i}: the short-term relationship
β_1,…, β_6: the long-term relationship of the model
ε_t: the error term of the model

Finally, we apply the Toda and Yamamoto (1995) causality test to specify the causal links between the variables. This test maintains the long-run information in the model, unlike other tests that make no room for long-run information.[9] The procedure requires the determination of the optimal lag length (k) as well as the maximal order of integration (d_{max}) by using the information criteria statistics such as the Akaike information criterion (AIC) and Scwhartz's information criterion (SIC). Consequently, a level VAR with lag length ($d_{max} + k$) is tested by using the standard Wald test statistic.

3.3. *Data*

We used a monthly data between January 1999 and December 2019. Our basic variables are the price series of soft wheat, durum wheat, corn and barley (USD/ton), exchange rate (MAD/USD), grain imports (MAD/ton), tax revenues (MDh) and foreign exchange reserves (USD).

The prices of corn (Gulf) and wheat (FCW1) were obtained from the International Cereals Council (CIC) and France Agrimer while durum wheat and barley data were collected from the World Bank Index Mundi database. The exchange rate and the foreign currency reserves were obtained from the International Financial Statistics of the International Monetary Fund (IMF). The monthly cereal import bill was obtained from the "*Office des Changes*" while the tax revenue was collected from the public finance statistics published by the General Treasury which mainly include revenue from customs duties and import value added tax (VAT).

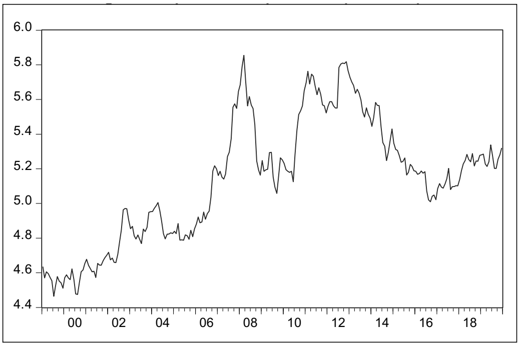
4. Results and discussion

Figure 3 depicts the imported cereal price series between 1999 and 2019. It shows the presence of significant price fluctuations characterized by bullish trends (uptrends) followed by bearish trends (2007-2009, 2011-2013 and 2016-2018) and, consequently, indicates that this series is volatile.

4.1. *Estimation of the price volatility function*

Following the Box-Jenkins Approach, we began by analyzing the correlogram of the differentiated series. The analysis of the simple and partial autocorrelation functions (Annex 2) shows that this is not a random walk because the critical probabilities are below 0.05. Therefore, we have a memory process that can be represented in the ARMA process class. Then, twenty-five combinations of ARMA (p, q) were computed. The comparison of the different processes using the SIC and the AIC indicates that ARIMA (0,1,1) is the best fit to explain the price series.

Figure 3 - Imported cereal price series (1999-2019).



Source: Produced by the author.

[8] We transformed all variables into logarithms to reduce variability and magnitude effects.
[9] Since the first differences and pre-whitening procedures are applied.

Table 1 - Unit root testing.

Variables	*ADF*	*PP*	*Integration*
Vol_ex	-1,11(0,24)	-1,07 (0,26)	I(1)
Vol_ex(-1)	-18,28 (0,000***)	-18,15 (0,000***)	
Vol_gr	-0,83(0,35)	-0,83(0,36)	I(1)
Vol_gr(-1)	-15,02 (0,000***)	-15,01 (0,000***)	
Res	2,36 (0,996)	2,03 (0,990)	I(1)
Res(-1)	-14,05 (0,000***)	-14,32 (0,000***)	
Fisc	4,08 (1,000)	2,45 (0,997)	I(1)
Fisc(-1)	-15,08 (0,000***)	-94,96 (0,000***)	
imp	2,82 (0,999)	3,22 (0,999)	I(1)
Imp(-1)	-5,69 (0,000***)	-30,03 (0,000***)	
Int	-12,84 (0,000***)	-12,76 (0,000***)	I(0)

**** indicates rejection of the null hypothesis of associated statistical tests at the 1% level of significance.*
() p-value. Source: Produced by the author.

The application of the Ljung-Box test confirms the null hypothesis of no autocorrelation of the residuals. The analysis of the normality of residuals with the Jarque-Bera indicator shows that normality is rejected (Annex 3). The testing for ARCH effects indicates that the null hypothesis of no ARCH effects is rejected (Annex 4).

In the next step, four models were selected and tested: ARCH, GARCH, TGARCH and EGARCH. Analysis and comparison of the estimation results of these models leads to the selection of the T-GARCH (1,1) (Annex 5). The same steps were followed to estimate the exchange rate volatility function and led to select the GARCH (1,1) model as the most appropriate.

4.2. *Unit root testing*

In order to use stationary series in our estimations, we apply Augmented Dickey Fuller (ADF) and Phillips-Perron (PP) unit root tests.[10] Results suggest first-order integration (I(1)) in all the series at 1% level with the exception of the variable "*Int*" which is I(0) (Table 1).

4.3. *Cointegrating analysis*

As emphasized by Pesaran *et al*. (2001), ARDL bound testing is based on F-statistic which assumes under the null hypothesis that lagged levels of the variables are jointly equal to zero. In order to obtain the optimal model, we employ Schwarz Bayesian Criteria (SBC) for selecting the optimum number of lags on each variable. The results show that the optimal lag for the variables in our model is ARDL (12,0,0,1,0,1) (Annex 6).

Then, several tests were conducted to verify the estimated model. These test include the Breusch-Godfrey correlation test, the heteroskedasticity and error normality test, the model stability test (CUSUM and CUSUMQ) proposed by Brown *et al*., (1975) and the error autocorrelation test. The results show that the selected model is good and explain 97% of the evolution of the economic bill (Annex 7). In addition, the correlogram indicates the absence of autocorrelation of the residuals, and the plot of CUSUM and CUSUM of squares statistics stay within the critical bounds, thus providing strong evidence of parameters stability in our model.

Later, we apply the bound test to assess the existence of the long-term relationship between the variables. The results indicate that the F-statistic (10.83) is greater than the upper bounds at all levels of significance (4.21) (Table 2). The hypothesis of no cointegration is rejected, and we conclude that there is cointegration between

[10] When ADF and PP tests are performed on the series, we do not find structural breaks.

Table 2 - Bound test results.

F- Statistic	*Lower Bound Value I(0)*	*Upper Bound Value I(1)*	*Critical value*
10,83	2.82	4.21	1%
	2.14	3.34	5%
	1.81	2.93	10%

Source: Produced by the author.

Table 3 - Estimation results of the long-run relationship.

Variable	*Coefficient*	*t-satistic*	*Probability*
Res	0,28	3,07	0,002***
Fisc	0,74	6,77	0,000***
Vol_gr	0,07	2,66	0,008***
Vol_ex	-0,28	-4,71	0,000***
Int	16,70	2,63	0,009***

**** Significant at 1%. Source: Produced by the author.*

invoice import, tax revenues, foreign exchange reserves, price and exchange volatility.

The analysis of the coefficients of the long-run cointegrating relationship (Table 3) shows that all variables are statistically significant at 1%. That is, foreign exchange reserves have a less than proportional positive effect on the import bill. As a result, a 1% rise in the foreign exchange reserves leads to a rise of 0.28% in import bill. This implies that the country requires more reserves in the long term to deal with growing import demand. Similarly, tax revenues increase proportionally with the import bill: a 1% increase in tax revenues increases the import bill by 0.74%. Customs revenues and import VAT increase with the increase of imports.

In addition, a 1% increase in price volatility and the volatility-import level (Int) leads respectively to a rise of 0.07% and 16.7% in the import bill. This implies that when quantity of imported grain decreases in good harvest years, price volatility no longer has a significant effect on the import bill. Specifically, cereal price volatility increases as grain imports increase. However, exchange rate volatility has a less than proportional negative effect.

In the next step, an error correction model (ECM) was computed to present the short-term dynamics that exist between import bill and other explanatory variables. As reported in Table 4, the coefficient of error correction term (ECT) is negative and statistically significant suggesting a quick adjustment process and, thus, implying that 100% of the disequilibrium from the previous month's shock adjusts back to equilibrium in the current month. However, unlike in the long run, foreign exchange reserves have a less than proportional negative effect on the import: a 1% decrease in foreign exchange reserves increases the import bill by 0.22% indicating that Moroccan imports put pressure on foreign exchange reserves in the short run. Import tax revenues bear a significant positive sign suggesting that a 1% rise in these revenues leads to a rise of 0.15% in import bill.

As for price volatility and the variable we created (Int), our results show that these variables exert a negative and significant effect (at different thresholds) in the short run. These effects reverse over time, indicating that the lagged value of these variables will have a positive impact on the next month's import bill. The time dimension is therefore an important variable to consider here. Over time, the effects of price volatility are mixed: it takes at least one month to expect to see the negative consequences of increased grain price volatility on the import bill. On the other hand, the effect of exchange rate volatility is negative and significant at 10%, indicating that a 1% increase in exchange rate volatility may lead to a 0.07% import bill decrease. One possible explanation for this surprising result

Table 4 - Estimation results of the short run relationship.

Variable	*Coefficient*	*t-statistic*	*Probability*
D(imp(-1))	0,48	2,66	0,008***
D(Res)	-0,22	-1,70	0,089*
D(fisc)	0,15	4,36	0,000***
D(fisc(-1)	0,19	5,29	0,000***
D(vol_gr)	-0,08	-2,02	0,044**
D(vol_gr(-1))	0,01	0,16	0,869
D(Vol_ex)	-0,07	-1,72	0,086*
D(Int)	-1,20	-2,24	0,026**
D(Int(-1))	2,28	2,64	0,009***
ECT(-1)	-1,00	-5,21	0,000***

****, **, * respectively significant at 1%, 5% et 10%. Source: Produced by the author.*

Table 5 - Results of Toda-Yamamoto causality test.

Variables dépendantes	*Variables explicatives ou causales*					
	Imp	*Res*	*Fisc*	*Vol_gr*	*Vol_ex*	*Int*
Imp		9,41***	3,47	4,26	7,20**	12,34**
Res	0,84		7,20**	0,85	1,14	1,45
Fisc	85,10***	3,42		0,70	0,08	0,05
Vol_gr	0,98	0,48	1,13		2,31	643,89***
Vol_ex	0,71	2,34	0,69	3,15		2,96
Int	0,75	2,17	2,02	5,28*	3,12	

****, **, * respectively significants at 1%, 5% et 10%. Source: Produced by the author.*

is that importers proceed, generally, to the exchange and hedging operation even before the closing of the transaction with suppliers.

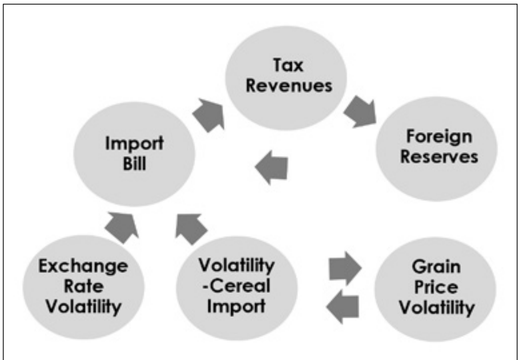
4.4. *Causality test*

As mentioned earlier, the Toda and Yamamoto (1995) test is applied to specify the causal links between the variables (Table 5).

The results show the following relationships (Figure 4):

- Bidirectional causality between grain price volatility and the volatility-import variable. The domestic level of import demand for cereals has an impact on price volatility. This latter influences the domestic import demand component. Such an effect could be observed in years of low domestic supply or during the Coronavirus pandemic, as the government suspends customs duties to encourage operators to import more and build up a safety stock.

Figure 4 - Causal links between the variables.



Source: Produced by the author.

- Unidirectional causalities:
- Foreign exchange reserves, exchange rate volatility and the volatility-import level variable impact the economic import bill.
- Import tax revenues influence the country's foreign exchange reserves and the import bill causes import tax revenues. Thus, the economic invoice of imports does no directly cause foreign exchange reserves because it was channeled through tax revenues.

5 Conclusion

This paper aims to evaluate the impact of grain price volatility on three macroeconomic indicators, namely import tax revenues (VAT and customs duties), foreign exchange reserves and the import bill for goods and services using monthly data between January 1999 and December 2019. We used: I) TGARCH (1,1) and GARCH (1,1) models to characterize the price and exchange rate volatility functions respectively; II) cointegration analysis and ARDL model to determine the long and short-run relationship between the variables; and III) Toda and Yamamoto's (1995) causality test to study the causal links between the different variables.

The results show that in the long-run, all variables are highly significant (probability = 1%). Foreign exchange reserves, import tax revenues, price volatility and the volatility-import levels variable all have a positive effect. Thus, a 1% increase in these variables amplifies the economic import bill by 0.28%, 0.74%, 0.07% and 16.7% respectively. It implies that an increase of the imports invoice is associated with an increased demand for foreign exchange reserves, increased grain price volatility and cash flow gain from customs import revenues. In addition, price volatility will have serious consequences for the balance of trade especially in years of low domestic supply. However, exchange rate volatility has a less than proportional negative effect on imports on the long run.

On the other hand, in contrast to the results for the long-run relationship, the effects of foreign exchange reserves, price volatility and the volatility-import level are negative and significant in the short term. As a result, a 1% rise in these three variables leads to a decrease in the import bill of 0.22%, 0.08% and 1.2% respectively. Our results also suggest that the effects of price volatility and the created variable are reversed over time. This indicates that the lagged value of these variables will have a positive impact on the next month's import bill. It takes at least one month to expect to see the negative consequences of increased grain price volatility on the import bill. Given this, the effect of import tax revenues remains positive and very significant in the short term since a 1% increase in these revenues increases the economic bill by 0.15%. In addition, the effect of exchange rate volatility on imports remains negative. This result does not corroborate those reported in similar studies (Vergil, 2002; Dönmez and Magrini, 2013; Ishimwe and Ngalawa, 2015) that suggest a significant impact of exchange rate volatility. The main reason for this finding is that grain importers proceed, generally, to the exchange and hedging operation even before the closing of a purchase agreement.

Furthermore, the results of the causality test indicate the presence of a bidirectional causality between grain price volatility and volatility-import level and unidirectional causalities between the other variables. Indeed, domestic demand for cereals on the international market influences price volatility and the import demand. Moreover, a one-way relation of causality from foreign exchange reserves, exchange rate volatility and volatility-import variable to the economic import bill was determined. This latter has an indirect impact on foreign exchange reserves through tax revenues. This is consistent with Ehrhart and Guerineau (2013) who found positive and significant effects of commodity prices volatility on tax revenues in developing countries.

Taking into account the results of this paper, Morocco is now more than ever called to reconsider the issue of grain food security as a strategic priority. Indeed, the country must increasingly work to secure its strategic stocks to confront the negative consequences of grain price volatility. This could occur through the implementation of concrete measures. Firstly, it is pointless to establish national strategies to ensure a high and sustainable local production (at both quantitative and qualitative level) irrespective of climate conditions. Secondly, it would be preferred to regularly explore the feasibility of undertaking a sovereign hedging program as was the case for oil in 2013 and butane in 2018. Thirdly, we believe that it would be more relevant to encourage the development of a large storage infrastructure (port infrastructure and storage silos) in order to optimize the unloading of the vessels, decrease demurrage costs and facilitate the reconstitution of strategic stocks.

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Annex 1 - Price index calculation steps

The import price index is calculated by the following formula:

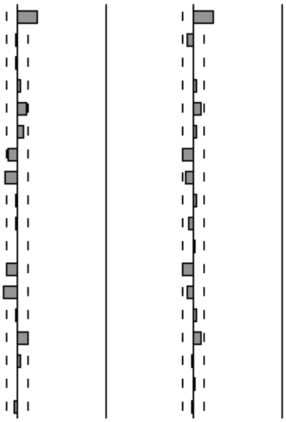
$$I_{i,t} = \prod_{c=1}^{4} p_{c,t}^{w_c}$$

Where $p_{c,t}$ is the international price of cereal c in month t. These prices are denominated in United States dollars and are expressed in CIF (Cost, Insurance and Freight).

The weight w_c is an average over the base period j (January 1999 to December 2019) of the share of imports of cereal c in total imports of the four cereals $w_c = {P_{jc}Q_{jc}}/{\sum_n P_{jn}Q_{jn}}$. The weight of each commodity is then held constant over time.

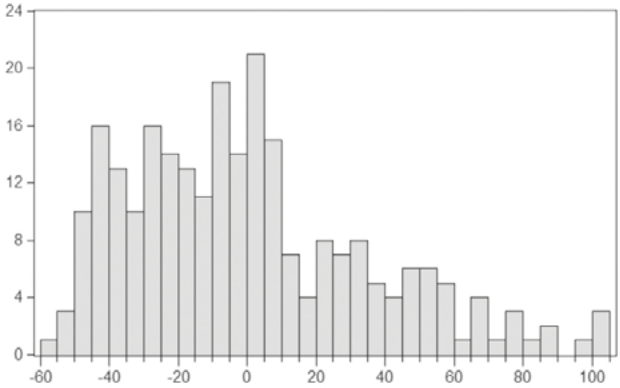
This index combines the prices of soft wheat, durum wheat, maize and barley and is similar to that calculated by Dehn (2000) and Ehrhart and Guerineau (2013). However, it differs from their indices in two respects. Firstly, it takes just the top four cereals imported by the country rather than all commodities. Secondly, it uses monthly data, which is more appropriate in the context of measuring volatility instead of annual or quarterly data.

Annex 2 - Correlograme



Autocorrelation	Partial Correlation		AC	PAC	Q-Stat	Prob
		1	0.228	0.228	13.175	0.000
		2	-0.013	-0.069	13.221	0.001
		3	-0.017	0.002	13.292	0.004
		4	0.032	0.036	13.551	0.009
		5	0.100	0.088	16.146	0.006
		6	0.073	0.033	17.533	0.008
		7	-0.111	-0.137	20.733	0.004
		8	-0.150	-0.094	26.606	0.001
		9	-0.015	0.033	26.663	0.002
		10	-0.021	-0.049	26.781	0.003
		11	0.006	0.016	26.791	0.005
		12	-0.136	-0.133	31.708	0.002
		13	-0.165	-0.078	38.950	0.000
		14	-0.020	0.034	39.055	0.000
		15	0.116	0.091	42.692	0.000
		16	0.031	-0.023	42.955	0.000
		17	-0.002	0.020	42.956	0.000
		18	-0.038	-0.015	43.352	0.001

Annex 3 - Normality test



Series: Residuals
Sample 1999M01 2019M12
Observations 252

Mean	0.084923
Median	-5.105743
Maximum	103.0172
Minimum	-56.58213
Std. Dev.	34.93266
Skewness	0.771101
Kurtosis	3.107328
Jarque-Bera	25.09399
Probability	0.000004

Annex 4 - ARCH test

Heteroskedasticity Test: ARCH

F-statistic	126.7511	Prob. F(1,249)	0.0000
Obs*R-squared	84.66916	Prob. Chi-Square(1)	0.0000

Test Equation:
Dependent Variable: RESID^2
Method: Least Squares
Date: 08/22/20 Time: 18:19
Sample (adjusted): 1999M02 2019M12
Included observations: 251 after adjustments

Variable	Coefficient	Std. Error	t-Statistic	Prob.
C	4063.530	633.4186	6.415236	0.0000
RESID^2(-1)	0.580618	0.051572	11.25838	0.0000

Annex 5 - Estimation results of ARCH, GARCH and TGARCH models[11]

Model	*R-sq*	*AIC*	*SIC*	*LB*[12]	*LB(2)*
ARCH(1)	0,0508	7,7138	7,7561	36,192 (0,460)	16,954 (0,997)
GARCH(1,1)	0,0488	7,4422	7,4986	24,998 (0,916)	15,173 (0,999)
T-GARCH(1,1)	0,0412	7,3590	7,4294	25,926 (0,892)	22,392 (0,963)

[11] EGARCH (1,1) is not accepted because it has two insignificant coefficients (C(3) and C(4)). Our selected model is TGARCH (1,1) because it has the lowest AIC and SIC.

[12] Ljung-Box Q-statistic of residuals (LB). The values in parentheses represent the probabilities (at 36 lags).

Annex 6 - Optimal lag structure and Estimation results of ARDL Model

Figure 5 - Optimal lag structure.

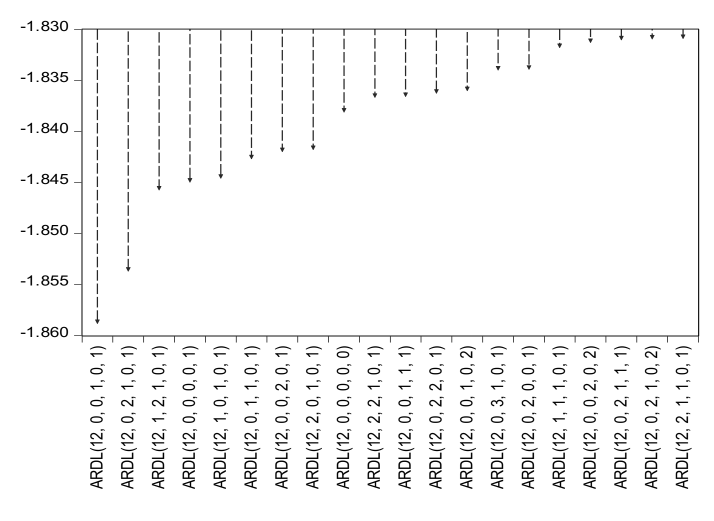


Figure 6 - Estimation results of ARDL(12, 0, 0, 1, 0, 1).

Selected Model: ARDL(12, 0, 0, 1, 0, 1)

Variable	Coefficient	Std. Error	t-Statistic	Prob.*
LIMP(-1)	0.306368	0.063217	4.846249	0.0000
LIMP(-2)	0.135239	0.066255	2.041186	0.0424
LIMP(-3)	0.124821	0.066878	1.866405	0.0633
LIMP(-4)	-0.099538	0.066174	-1.504185	0.1340
LIMP(-5)	0.197249	0.065528	3.010156	0.0029
LIMP(-6)	0.027896	0.066686	0.418313	0.6761
LIMP(-7)	-0.062603	0.066508	-0.941283	0.3476
LIMP(-8)	-0.194272	0.064275	-3.022522	0.0028
LIMP(-9)	0.096670	0.065296	1.480497	0.1402
LIMP(-10)	-0.027866	0.065158	-0.427673	0.6693
LIMP(-11)	0.096919	0.063720	1.521016	0.1297
LIMP(-12)	0.210217	0.058702	3.581058	0.0004
LFISC	0.139967	0.032666	4.284750	0.0000
LRES	0.052140	0.019727	2.643100	0.0088
VOL_GR	-0.101479	0.040055	-2.533495	0.0120
VOL_GR(-1)	0.115061	0.040040	2.873652	0.0045
VOL_EX	-0.052242	0.012934	-4.039189	0.0001
INT	-0.445309	0.510898	-0.871619	0.3844
INT(-1)	3.600928	0.936503	3.845077	0.0002

R-squared	0.973942	Mean dependent var	10.00627
Adjusted R-squared	0.971820	S.D. dependent var	0.477410
S.E. of regression	0.080143	Akaike info criterion	-2.134156

Annex 7 - Model diagnostic test results

Hypothesis	*Test*	*Value*	*Probability*
Autocorrelation	Breush-Godfrey	0,01	0,93
Heteroscedasticity	Arch-Test	1,01	0,31
Normality	Jarque-Bera	4,69	0,10
Specification	Ramsey	0,56	0,58

Figure 7 - Plot of CUSUM and CUSUM of squares Tests.

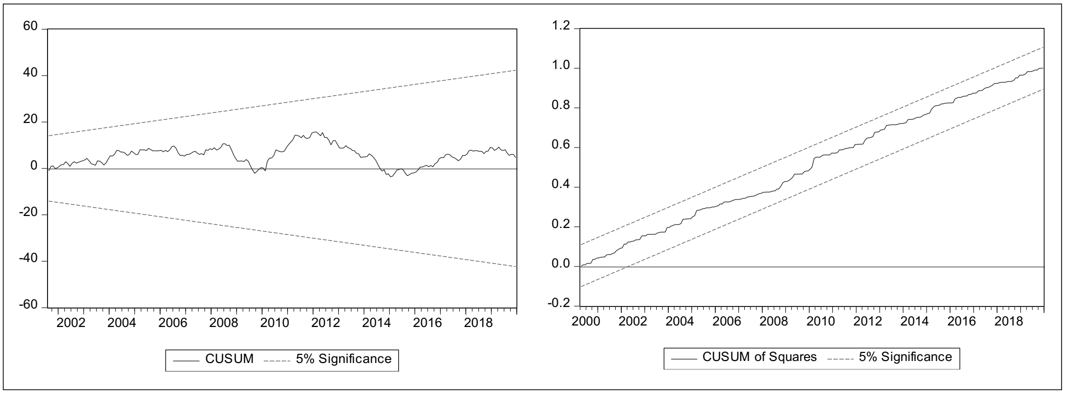


Figure 8 - Correlogram of residuals.

Autocorrelation	Partial Correlation		AC	PAC	Q-Stat	Prob*
		1	-0.003	-0.003	0.0015	0.969
		2	0.079	0.079	1.5392	0.463
		3	0.018	0.019	1.6195	0.655
		4	0.061	0.055	2.5314	0.639
		5	0.001	-0.001	2.5317	0.772
		6	-0.043	-0.053	2.9950	0.009
		7	0.033	0.036	3.2629	0.860
		8	-0.018	-0.014	3.3409	0.911
		9	0.071	0.079	4.6095	0.867
		10	0.031	0.042	4.8507	0.901
		11	0.041	0.036	5.2769	0.917
		12	0.150	0.145	11.013	0.528
		13	-0.068	-0.089	12.206	0.511
		14	0.033	0.000	12.490	0.567
		15	-0.112	-0.112	15.744	0.399
		16	-0.107	-0.127	18.738	0.282
		17	0.017	0.053	10.010	0.340
		18	-0.089	-0.060	20.870	0.286
		19	-0.108	-0.096	23.954	0.198
		20	0.004	0.029	23.958	0.244
		21	-0.030	-0.060	24.202	0.283
		22	0.006	0.002	24.213	0.336