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**Article**

# Technological progress and finance: The effects of digitization on Brazilian banking fees

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# Technological progress and finance: The effects of digitization on Brazilian banking fees

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ABSTRACT

Recent technological advancements and the digitization of most financial services have transformed the way by which individuals deal with their financial transactions. It has also affected the focus of financial institutions' investments, as well as the sectors' market dynamic, with the increasing number of start-ups and fintechs, which tends to lead to higher competition. Considering this scenario, our study explores the impacts that mobile and home banking had on bank fees charged to Brazilian consumers, from 2012 to 2019. As our theoretical framework, we use a network-city model, which is an evolution of the seminal linear and circular city models, while the econometric analysis is made using a time series model. Our findings show that bank fees were negatively affected by the digital banking platforms. In summary, the results indicate that the decrease on transport costs, caused by financial digitization and the widespread use of mobile and home banking, has reduced the price of bank fees to customers in the studied period.

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## 1. Introduction

The Brazilian banking sector continues to be characterized as an oligopolistic market structure, as the last decades have been marked by privatizations and many mergers and acquisitions (Hordones, 2019), despite the recent emergence of digital banks and the increase in the number of fintechs (Barbosa, 2018). Bittencourt et al. (2015) report that for years now, Brazilian financial institutions have been (following a world trend) undergoing merger processes – which intensified with the 2008 subprime crisis – and this opens debate about the effects on financial stability. According to the Central Bank of Brazil (BACEN, 2020), the Concentration Ratio of the five biggest Brazilian banks (CR5), in terms of assets, was of 81% in December 2019 and the normalized Herfindahl-Hirschman Index (HHIn) was 0.1367, making Brazil one of the countries with the most concentrated banking sectors in the world.

The fintech start-ups ecosystem has grown in recent years — BACEN (2020) reports that, in June 2019 there were 604 fintechs, representing a 33% growth since the last report in 2018. Thus, there is space for discussion about the impacts of digital companies on the competition of the banking sector over time. The digitization of this business model reduces marginal

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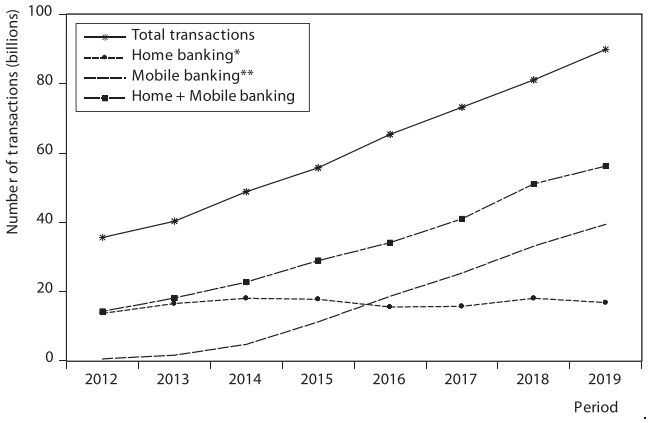


**Fig. 1.** Number of bank transactions between 2012 and 2019 in Brazil. *Home banking includes all transactions via the financial institution's website. **Mobile banking includes transactions using a smartphone/mobile applicationor software.
Source: the authors, with data from FEBRABAN (2020).

operation and switching costs, which makes it easier for new firms to enter the market. The digital business model is also characterized by strong sales and earnings growth, both driven by an innovation environment, which enables these companies to accelerate the development of new projects and investments for the market (Dapp, 2015; Coetzee, 2018).

In Brazil, digital banking channels already account for 25% of the financial transaction operations (raising from 7.5 billion transactions in 2018 to 8.8 billion transactions in 2019), and 86% of the non-financial operations, as reported by FEBRABAN (2020) in a research with 22 financial institutions which represent 90% of the sector's assets in Brazil. Brazilian banks' expenses in new technologies reached R$ 24.6 billion in 2019 — a 24% growth over the previous year, with an increase of 48% in investments (FEBRABAN, 2020). In light of this, discussions about the relevance and future of the traditional bank branches gain space. Furthermore, amidst the new market reality, it is not possible to discuss competition in the banking sector without mentions to the emergence and consolidation of digital banks and fintechs in the country, which follows a worldwide trend.

As it is seen from Fig. 1, total banking transactions grew year by year, and so did transactions via digital channels. We can also see that, around 2016, transactions made using mobile applications have surpassed the ones made through regular home banking. It is very likely that this fact is one of the results of the expansion and popularization of smartphones in society. Together, the two channels accounted for more than half of the bank transactions made in Brazil in 2019, reaching 63% of the total transactions number.

Considering this scenario, one important question to be answered is whether bank charges and fees have been affected by the institution and expansion of digital platforms (e-banking). We see two possible channels of effect: first, the high investments in the developments of new technologies (R&D, software, applications and safety measures) could impact the profit margins of financial institutions, leading to a transfer of costs to consumers, which would potentially lead to tariffs increasing (throughout the remaining of the paper, the terms "fees" and "tariffs" are used interchangeably).

On the other hand, the high investments in technology could make it possible for financial institutions to reduce expenses such as physical branches, infrastructure and headcount, which would allow them to reduce their fees, seeking to be more competitive. In addition, the dissemination of remote services could cause a decrease in horizontal (locational) differentiation among banks, making them more homogeneous, which would also contribute to lower fees (Degryse, 1996; Veiga and Oliveira, 2006). The fundamental argument is that, due to market concentration and the fact that only few banks hold high market power, bank tariffs are set in collusion, and e-baking could reduce the localization effect and informational asymmetry in the sector, generating pressures on banks to reduce their fees.

Therefore, this paper explores the effects that the recent popularization of e-finance had on the bank fees charged to Brazilian customers, from an empirical analysis between the years of 2012 and 2019. A similar study was performed by Veiga and Oliveira (2006), in which the authors, analyzing data from 1996 to 2002, found evidence that an increase in the number of ATMs available in Brazil lowered the bank fees charged to consumers. Therefore, we build on their work to expand and update this analysis, taking into account the recent technological advancements on e-banking, mainly the widespread use of home and mobile banking, which have already surpassed the use of ATMs. Our study relies on Wang and Wang's (2018) network-city model — an evolution of the seminal works of Hotelling (1929) and Salop (1979) — to measure the spatial aspects of the Brazilian banking sector structure.

The paper is structured as follows: in Section 2 we discuss the effects of e-banking on competition, and we also present the network-city model. Next, Section 3 contains the details of our methodology, including the data we used and our econometric approach. Following, on Section 4 we present our main results. We then complete the paper with Section 5, in which we draw our conclusions and final thoughts about the study.

## 2. Theory

### 2.1. The banking sector and competition in Brazil

The way by which society — individuals and companies — deals with their money has evolved constantly, along with the technological progress achieved during the last decades. The history of currency dates back over 4000 years, beginning with barter, then going through the creation of paper money, the later emergency of the first credit cards in 1920, until finally, in the 1990 s, the rise and popularization of internet banking (Galbraith, 1997). Technological progress has revolutionized financial markets, enabling banks to offer a growing range of automated products and services to their customers (Veiga and Oliveira, 2006). From ATM and ATM cards to the introduction of the internet and digital banks, the organization of the banking sector, as well as the relationship between banks and society, has gone through profound changes.

The analysis performed in this study follows Veiga and Oliveira (2006) definition of the Brazilian banking sector as an oligopoly, in which consumers can perceive differences between the products of different firms — that is, there is product differentiation and, as a result, firms have some level or market power. With the technological advancements that led to the expansion and popularization of mobile and internet banking, market competition on the banking sector has increased fiercely. Information about each bank's products, services, costs and benefits has become easily accessible, which increases the range of information possessed by consumers when choosing a bank.

In front of this evolution, discussions about the sector's organization are brought up. Most of the existing literature on the Brazilian economy tends to conclude that the banking sector is competitive, despite its high levels of concentration (Araújo et al., 2007, 2006; Lucinda, 2010; Nakane, 2001). The Central Bank of Brazil (BACEN, 2020), based on the Lerner and Boone indicators, argues that there was an upward trend in the degree of competition in the banking sector between 2000 and 2019. These conclusions, however, are not definitive, as they are challenged by other studies, such as the ones of Tabak et al., (2015), Barbosa et al., (2015), and Divino and Silva (2017), who find evidence suggesting that the Brazilian banking sector might actually not be competitive.

Despite these divergences regarding competition levels on the Brazilian banking sector, there is consensus that the emergence of technological innovations requires financial institutions to change their strategies and adapt their ways of providing services, in addition to building a more dynamic and competitive market (Bittencourt et al., 2015; Gomes et al., 2017; Coetzee, 2018). The way banks capitalize new opportunities derived from technological advancements — and deal with the potential threats they pose - will be critical to their survival in the future (Coetzee, 2018).

### 2.2. Theoretical framework

Initially, in order to show the effects of e-finance on competition, we will consider a network-city spatial model of competition, based on a Bertrand oligopoly, and developed by Wang and Wang (2018). This model can be used to analyze various issues related to industrial organization, such as equilibrium, diversification and mergers. Table 1 provides a brief review of the literature on this locational competition approach.

The literature shows that a business' location is an important factor in consumer choice decisions. However, over the years, advancements in transportation and technology have changed the effects of the location factor. The choice for Wang and Wang's (2018) model is based on its contemporaneity and the fact that it is a good representation of a real competitive setting. Also, their model relaxes some of the restrictions of more traditional competition models such as Hotelling's (1929) and Salop's (1979), allowing for heterogeneous constant marginal costs of production, besides other innovations. In their model (Fig. 2), firms are located at the vertices of the network and compete directly and simultaneously, while consumers are located along the links that connect the vertices and incur transportation costs when deciding to buy from a particular firm.

As seen from Fig. 2, there are $n$ firms spatially located at the vertices of a $n$-vertex network (Fig. 2 illustrates a 6 vertices network), with one firm per vertex. Every vertex is connected to the other by a link (called a linear city) of length 1. Thus, there is a total of $n(n-1)/2$ links or linear cities. Consumers are uniformly distributed, with the density at each link given by $f$. Firm $i$ has a constant marginal cost of production denoted by $c_i$, and sets price $p_i$, $i = 1, 2, ...,n$. Each consumer demands either 0 or 1 unity of the firm's product, with willingness to pay denoted by $v$, and transportation costs per unit of distance defined as $t$. The demand for firm's product $j$ comes from the $n-1$ links connected to firm $j$. Following Wang and Wang (2018), firm $j$'s Nash equilibrium price will be:

**Table 1**
Literature evolution of the competition models for localization.

| Author | Date | Model |
|---|---|---|
| Hotelling | 1929 | Linear city |
| Salop | 1979 | Circular city |
| Von Ungern-Sternberg | 1991 | Pyramid |
| Chen and Riordan | 2007 | Spokes |
| Somaini and Einav | 2013 | Dynamic spatial |
| Wang and Wang | 2018 | Net-work city |

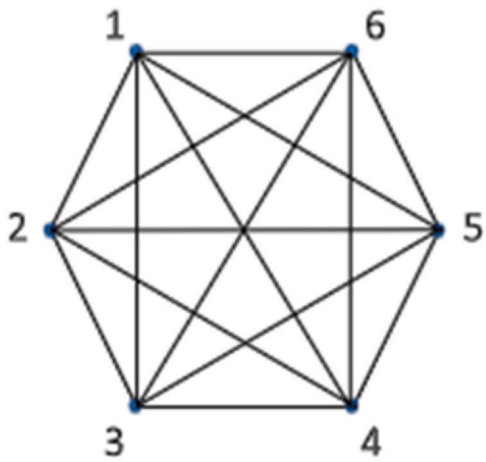


**Fig. 2.** Network-city model
Source: Wang and Wang (2018).

$$p_j = t + \frac{1}{2n-1}\left(\sum_{i=1}^{n} c_i + (n-1)c_j\right) \tag{1}$$

This way, regarding the transportation cost ($t$), the expected relationship is that an increase in $t$ would lead to higher prices ($p$), due to the horizontal (locational) differentiation factor. For consumers, the transportation costs can be perceived as a *disutility* derived from the need of physically going to a bank (whether to open a new account, switch from one bank to another, or withdrawing money) which is strongly reduced with the use of e-banking.

As for the marginal costs ($c$), we expect that the higher this variable is, the more banks would be willing to transfer such costs to consumers, potentially leading to an increase in tariffs ($p$). We can also see from the model that firm $j$ sets its prices considering not only its own costs, but taking into account the marginal costs of firm $i$ as well. As for the number of banks ($n$), since traditional models suppose homogenous marginal costs among firms, one would usually expect that prices would be independent of $n$. Even in Wang and Wang's (2018) model, when all firms are symmetric this condition would also be true. However, this model relaxes such hypothesis, allowing for any arbitrary number of firms with differentiated constant marginal costs. In this case, if new bank entrants have lower marginal costs than the ones of currently established banks, then an increase in the number of banks could lead to lower tariff prices.

## 3. Method

In this section we describe the empirical analysis methods we used to explore the effects of e-banking on bank fees for the period of 2012–2019. We follow the basic hypothesis that tariff prices would suffer a positive effect from higher expenses, transportation costs and sector concentration. Regarding transportation costs, we used the number of mobile and home banking transactions as an inverse for this variable, as the use of these technologies generates convenience and utility for consumers and decreases the horizontal (locational) differentiation of the banking sector. Since it is the inverse of the transportation cost, we could also put that tariff prices will be lower the higher the number of transactions made using mobile and home banking is.

For expenses, we considered all operational and administrative costs incurred by banks in their activities. Tariff prices should be higher the higher these expenses are. Finally, regarding the number of banks, we considered the concentration level in the market (HHI), and we expected that tariffs would be lower the lower sector concentration is, due to the increase in market competition.

### 3.1. Data

For this analysis, data from the following sources were used: Central Bank of Brazil (BACEN), which gathers information about all financial institutions authorized to operate in the country, and the Getúlio Vargas Foundation (FGV, 2020) database, to obtain the General Market Price Index (IGP-M) series, used to deflate the series in our model.

Regarding the data obtained from BACEN, it should be noted that we used the "Balance Sheet/General Balance Sheet" (BACEN, 2019a), which contains monthly data provided by every Brazilian financial institution. Among the consulted accounts, we highlight those with the codes "71700009" and "81700006", which are, respectively, the "Revenue from Services" and "Administrative Expenses" accounts – considering data availability, these were the accounts containing the most approximate data needed for our analysis. In addition, data derived from the Credit Guarantee Fund (BACEN, 2019b), which provides information on the totality of bank accounts in Brazil, were also used. Our variables are constructed over monthly data, covering

the period between January 2012 and December 2019, in a total of 96 observations. Details about the data accounted for in each of these databases, as well as descriptive statistics for all variables, are available in Appendix A.

### 3.2. Econometric model

Initially, the econometric model we built could be described as the following:

$$TARIFF_t = \beta_0 + \beta_1 EXP_t + \beta_2 MOBILE_HOME_t + \beta_3 CONCENTRATION_t + \varepsilon_t \tag{2}$$

Our dependent variable *TARIFF* corresponds to the element *p* described in Eq. (1) of the theorical model (Wang and Wang, 2018). To create this variable, we used the monthly[2] sum of total revenues originated from services rendered by all financial institutions, which are registered in BACEN's account number "71700009".[3] The variable was deflated by the IGP-M index and divided by the number of active bank accounts in each month,[4] as a way to calculate the average revenue per account.

Therefore, the equation for calculating *TARIFF* is:

$$TARIFF_t = \frac{\sum_{i=1}^{n} Revenue\ from\ Services_{t,i}}{\sum_{i=1}^{n} Accounts\ in\ the\ month_{t,i}} \tag{3}$$

in which *i* represents the sum of revenue from services for each bank *i*, and *t* represents the time period (month). Fig. 3 shows the evolution of *TARIFF* over the analyzed period. The average tariff price of the whole sample is R$ 20.18, although there is a lot of volatility in this amount. However, over the period, there is a clear downward trend in the overall average price of bank tariffs.

The *EXP* variable represents what are considered the marginal costs *(c)* for each financial institution *i* in the theoretical model of Eq. (1). Therefore, as mentioned at the beginning of Section 3, we expected that higher expenses would result in higher fees, hence, an increase in this variable should result in a positive effect on the dependent variable *TARIFF*.

For this variable, we used the values registered in account number "81700006", which represents the administrative expenses from the financial institutions. It is in this account that banks record expenses such as employee costs and rents. Total administrative expenses were added monthly for all financial institutions.[5] These values were deflated by the IGP-M index and applied in the form of billions of *Reais (R$)*. Fig. 4 shows the evolution of expenses in the studied period. Bank expenses are quite volatile and seem to not have a predictable trend, although there is a possible seasonality at the end of each year.

Regarding the transportation cost variable (*t* in the theoretical model), we used the total number of mobile and home/internet banking transactions (in billions) as a *proxy* representing the inverse of the transportation cost, and this variable was named as *MOBILE_HOME*. The main idea is that customers, when conducting bank transactions via mobile or home/internet banking, do not need to travel to their banking institution, that is, the use of these technological means serves as the inverse of the transportation cost for customers. A similar methodology is applied by Veiga and Oliveira (2006), however, the authors use the number of ATMs and electronic only branches as the inverse of the transportation cost.

In their Retail and Card Payment Statistics database, BACEN (2019c) reports data on the transactions made via *Home banking* and *Mobile banking,* which can be seen in Table 2. There is a growth on the use of these modalities over time, just as expected, given the expansion and widespread use of internet and smartphones seen in society, as well as the strong investments made by banks in new technologies that facilitate consumer access to their products and services. Thus, to build the *MOBILE_HOME* variable, we summed all the year-to-year number of transactions in mobile and home banking, with or without actual financial transfer, and these numbers were converted into monthly data.[6]

Finally, to represent the element *n* in the theoretical model, we created the *CONCENTRATION* variable. Historically speaking, the Brazilian finance sector has always been well concentrated, with most of market power held by four to five banks, which has, until recently, limited the entrance of new, smaller institutions - and, consequently, has limited the number of banks. Thus, we chose to use a measure of concentration instead of the number of banks in our model, as we believe it is a more realistic approximation for the Brazilian case.

As our measure of concentration, we used the Hirshman-Herfindahl Index (HHI), which is defined by the sum of the squares of each banks' market share. The HHI is the main index used by regulators to measure the degree of concentration

[2] The amounts are reported in a cumulatively way for all semesters, therefore, to reach the monthly amount, we subtracted the values of subsequent months within each semester.

[3] It is in this account that the financial institutions record several of their revenues, including those from bank tariffs or fees - but not exclusively. It would be more accurate to control only for the revenues that actually originate from tariffs, but this isolated data is not available.

[4] It is important to note that the data regarding the number of accounts are only available in a semiannual basis and, therefore, their semiannual amounts were replicated for all months that are part of that due semester.

[5] This variable is also cumulatively reported for each semester, so in order to reach the monthly amount we have subtracted the amounts registered for subsequent month within each semester.

[6] As the data is published on an annual basis, we used its average monthly growth to simulate the number of transactions in each month, so that these monthly numbers evolved in the same proportion as the annual growth, and that their sum for each year was equal to the total number of transactions published by BACEN for every year.

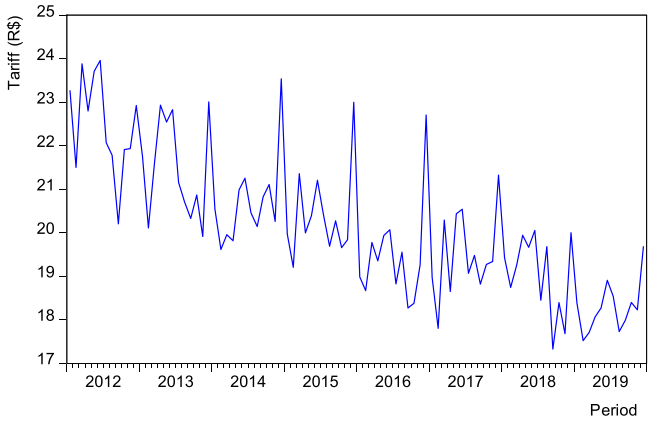


**Fig. 3.** Tariffs charged to consumers for bank services from 2012-2019.
Source: the authors, with data from BACEN (2019a, 2019b).

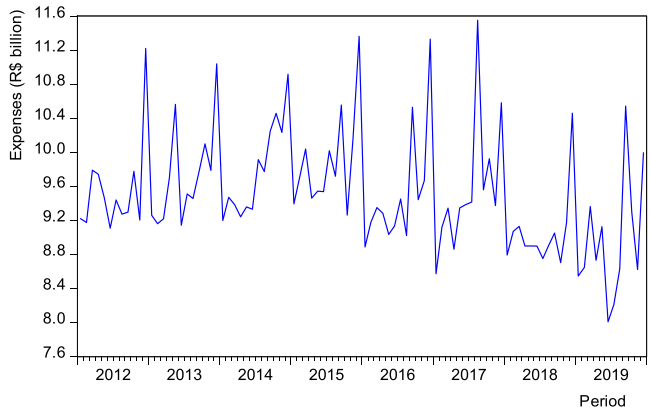


**Fig. 4.** Total bank expenses from 2012–2019.
Source: the authors, with data from BACEN (2019a, 2019b).

**Table 2**
Use of mobile and home banking in Brazil (billions of transactions).

| Year | *Home* | *Mobile* | Total |
|---|---|---|---|
| 2012 | 15.559 | 0.821 | 16.380 |
| 2013 | 17.740 | 2.290 | 20.030 |
| 2014 | 19.466 | 5.143 | 24.609 |
| 2015 | 19.726 | 10.799 | 30.525 |
| 2016 | 20.274 | 16.725 | 36.999 |
| 2017 | 20.605 | 24.531 | 45.136 |
| 2018 | 21.895 | 29.065 | 50.960 |
| 2019 | 22.870 | 33.916 | 56.786 |

Source: the authors, with data from BACEN (2019a, 2019b).

of a market (Hordones, 2019) and is well spread and used in the Industrial Organization literature. The equation which results on the HHI is:

$$HHI_t = \sum_{i=1}^{n} \left( \frac{X_{i,t}}{X_t} \right)^2 . \, 10000 \tag{4}$$

in which $X_{i,t}$ represents the revenues for the institution $i$ in period $t$, and $X_t$ represents the sum of revenues for all the sample's institutions in period $t$. This index can be calculated based on several items, such as the number of accounts by each institution, the number of customers, the total revenues, among others. In this work we chose to calculate it based on the revenues, since they are part of the main focus in our analysis. Following economic intuition, we expected that *CONCENTRATION* would have a positive effect on *TARIFF* in our econometric model, since an increase in market power would tend to lead to higher prices.

**Table 3**
Statistical tests and treatments.

| Issue | Test | Result | Treatment |
|---|---|---|---|
| Causality | Granger Causality Test | There is causality among the variables | – |
| Stationarity | Augmented Dickey-Fuller (ADF) | All variables follow an I(1) order of integration | Cointegration test |
| | Phillips-Perron (PP) | | |
| Cointegration | Johansen | Trace and Max-eigenvalue indicate 3 cointegrations | – |
| | Residuals stationarity | Residuals are stationary | – |
| Residual normality | Jarque-Bera | Residuals are normally distributed | – |
| | Kolmogorov-Smirnova | | – |
| Residual heteroskedasticity | Breusch-Pagan-Godfrey | Residuals are homoscedastic | – |
| | White | | – |
| Residuals autocorrelation | Breusch-Godfrey (LM) | There is autocorrelation in the residuals | HAC (Newey-West estimator) |
| | Durbin-Watson | | |

## 4. Results

### 4.1. Econometric tests and treatment

Before the analysis of the results, we conducted some statistical tests and treatments that confirmed the robustness of the model, as per Table 3.

First, to validate the existence of an effective connection between the theoretical model variables and the ones used to represent them in our econometric estimation, we employed the Granger Causality Test. We tested all the variables up to 12 lags and the results confirmed that the main variable of our analysis, *MOBILE_HOME,* indeed had causal effects on *TARIFF* in almost every lag. Considering that *MOBILE_HOME* was a *proxy* variable that we created to represent the inverse of transportation costs in the theoretical model, this result allowed us to move on with our estimation. The complete results can be consulted in Appendix B.

Then, since we are running a time series model, all the variables should be stationary. Therefore, we ran the Augmented Dickey-Fuller (ADF) and the Phillips-Perron (PP) tests, and the results (Table 4) indicated that none of the variables are stationary in level, since it was not possible to reject the null hypothesis in all cases (intercept, trend and intercept, and none) for at least one test. We then tested these variables in their first difference, and the results indicated that they follow a *I(1)* order of integration, i.e., they are stationary in their first difference.

As all variables are *I(1)*, the usual procedure would be to consider them in their first difference in the model. However, since our primary concern is on the long-term relationship of the variables, it is possible to conduct the regression at level, provided that the variables are cointegrated. In addition, to prove that the variables hold a long-term relationship, confirming cointegration is important in order to eliminate the risk of a spurious regression (Gujarati and Porter, 2011).

We applied the same stationarity tests on the generated residuals. If some time series generates stationary residuals — even if obtained from non-stationary variables — then one can infer that such series are cointegrated. The test on the residuals is in Table 5, and the results confirm that the residuals of the model are, in fact, stationary.

Along with the stationarity tests of the residuals, we have also run Johansen's cointegration test and the results allow us to not reject the existence of cointegration. Table 6 indicates a considerable number of cointegrating relationships in every

**Table 4**
Stationarity tests at the level and in their first difference.

| Variable | Test | t-statistic | | | |
|---|---|---|---|---|---|
| | | Intercept | Trend and intercept | None | I |
| TARIFF | ADF | -1.187193 | -2.125115 | -2.896229 * ** | |
| | PP | -4.922697 * ** | -9.151855 * ** | -1.013938 | |
| EXP | ADF | -8.421845 * ** | -9.030533 * ** | -0.97302 | |
| | PP | -8.714025 * ** | -9.108926 * ** | 0.094499 | |
| MOBILE_HOME | ADF | 1.975465 | -2.304601 | 13.45538 | |
| | PP | 1.95649 | -2.304025 | 11.06137 | |
| CONCENTRATION | ADF | -5.441904 * ** | -6.181532 * ** | 0.423458 | |
| | PP | -5.367679 * ** | -6.187961 * ** | -0.315754 | |
| ΔTARIFF | ADF | -5.945719 * ** | -5.963313 * ** | -5.021147 * ** | I(1) |
| | PP | -66.48612 * ** | -75.75660 * ** | -34.06749 * ** | |
| ΔEXP | ADF | -7.706156 * ** | -7.812725 * ** | -7.649957 * ** | I(1) |
| | PP | -35.72317 * ** | -35.82519 * ** | -36.07599 * ** | |
| ΔMOBILE_HOME | ADF | -9.029329 * ** | -9.423471 * ** | 0.243982 | I(1) |
| | PP | -9.053849 * ** | -9.423449 * ** | -3.73757 * ** | |
| ΔCONCENTRATION | ADF | -9.315210 * ** | -9.457484 * ** | -9.349294 * ** | I(1) |
| | PP | -24.99928 * ** | -29.13591 * ** | -25.59527 * ** | |

Note: *, **, *** indicates rejection of the null hypothesis at 10%, 5% and 1% respectively

**Table 5**
Stationarity tests of the residuals.

| Variable | Test | t-statistic | | | |
|---|---|---|---|---|---|
| | | Intercept | Trend and intercept | None | I |
| RESID | ADF | -7.608939 * ** | -7.563781 * ** | -7.64896 * ** | I(0) |
| | PP | -7.774054 * ** | -7.73367 * ** | -7.812884 * ** | |

Note: *, **, *** indicates rejection of the null hypothesis at 10%, 5% and 1% respectively.

**Table 6**
Johansen cointegration test (all specifications).

| **D**ata Trend: | None | None | Linear | Linear | Quadratic |
|---|---|---|---|---|---|
| **Test Type** | No Intercept No Trend | Intercept No Trend | Intercept No Trend | Intercept Trend | Intercept Trend |
| Trace | 3 | 3 | 2 | 2 | 2 |
| Max-Eig | 3 | 3 | 2 | 2 | 2 |

Selected (0.05 level*) Number of Cointegrating Relations by Model
*Critical values based on MacKinnon et al. (1999)

specification. Meanwhile, for the specification considered in our model (with intercept and no trend) there are at least 3 cointegrating equations (Table 7). These results allow us to go on with the long-term relationship analysis of the series.

After confirming for cointegration, we checked the residuals for normality, heteroscedasticity, and serial autocorrelation. As seen in Table 8, the Jarque-Bera and Kolmogorov-Smirnova tests confirm the null hypothesis of normality, evidencing that the residuals are normally distributed.

Regarding heteroscedasticity, we applied White's test, which indicated that the residuals are homoscedastic, since the statistical results presented on Table 9 allow us to not reject the null hypothesis of homoscedasticity.

As for autocorrelation on the residuals (Table 10), we ran a Durbin-Watson test, and the d value was 1.475044, which already indicates the presence of autocorrelation, since, considering a 100 observations sample with 3 dependent variables, the d value found was below the critical value that would negate autocorrelation. In addition, we also ran the Breusch-Godfrey's — also

**Table 7**
Johansen cointegration test (with intercept and no trend).

Johansen Cointegration Test
Trend assumption: No deterministic trend (restricted constant)
**Unrestricted Cointegration Rank Test (Trace)**

| Hypothesized No. Of CE(s) | Eigenvalue | Trace Statistic | 0.05 Critical Value | Prob.[b] |
|---|---|---|---|---|
| None[a] | 0.328391 | 97.44866 | 54.07904 | 0 |
| At most 1[a] | 0.307943 | 60.42729 | 35.19275 | 0 |
| At most 2[a] | 0.186116 | 26.19515 | 20.26184 | 0.0067 |
| At most 3 | 0.072934 | 7.042984 | 9.164546 | 0.1242 |

Trace test indicates 3 cointegrating eqn(s) at the 0.05 level
**Unrestricted Cointegration Rank Test (Maximum Eigenvalue)**

| Hypothesized No. Of CE(s) | Eigenvalue | Max-Eigen Statistic | 0.05 Critical Value | Prob.[b] |
|---|---|---|---|---|
| None[a] | 0.328391 | 37.02137 | 28.58808 | 0.0033 |
| At most 1[a] | 0.307943 | 34.23214 | 22.29962 | 0.0007 |
| At most 2[a] | 0.186116 | 19.15217 | 15.8921 | 0.0148 |
| At most 3 | 0.072934 | 7.042984 | 9.164546 | 0.1242 |

Max-eigenvalue test indicates 3 cointegrating eqn(s) at the 0.05 level
[a] denotes rejection of the hypothesis at the 0.05 level
[b] (MacKinnon et al., 1999) p-values

**Table 8**
Residual normality tests.

| Test | Statistic | Probability |
|---|---|---|
| Jarque-Bera | 3.623499 | 0.163368 |
| Kolmogorov-Smirnova | 0.075112 | 0.200000 |

Null Hypothesis: the data is normally distributed

**Table 9**
Heteroskedasticity tests.

| | | | |
|---|---|---|---|
| **Breusch-Pagan-Godfrey test** | | | |
| F-statistic | 0.920986 | Prob. F(3,92) | 0.434 |
| Obs*R-squared | 2.799026 | Prob. Chi-Square(3) | 0.4237 |
| Scaled explained SS | 2.087955 | Prob. Chi-Square(3) | 0.5544 |
| **White test** | | | |
| F-statistic | 1.359278 | Prob. F(9,86) | 0.2194 |
| Obs*R-squared | 11.95536 | Prob. Chi-Square(9) | 0.2158 |
| Scaled explained SS | 8.918192 | Prob. Chi-Square(9) | 0.4449 |

Null Hypothesis: variances for the errors are equal (homoscedasticity)

**Table 10**
Residuals serial correlation tests.

| | | | |
|---|---|---|---|
| **Breusch-Godfrey (LM) test** | | | |
| F-statistic | 6.164409 | Prob. F(1,91) | 0.0149 |
| Obs*R-squared | 6.090535 | Prob. Chi-Square(1) | 0.0136 |
| **Durbin-Watson test** | | | |
| Durbin-Watson stat | 1.475044 | | |
| dL | 1.602 | dU | 1.732 |

Null Hypothesis: There is no serial correlation in the residuals
Note: dL and dU calculated for n = 100 and k′ = 3

known as LM — test. The results were positive, at a 5% significance level, for the presence of autocorrelation, confirming the results of the Durbin-Watson test.

As seen, the model presented autocorrelation on the residuals, which means that any results could be biased. To avoid this problem, Gujarati and Porter (2011) suggest the correction of the standard errors using a procedure known as HAC (heteroscedasticity-and-autocorrelation-consistent) standard errors, also known as Newey-West standard errors. Thus, we used this procedure in order to correct the standard error for the presence of autocorrelation and the residual problems of the model.

### *4.2. Regression results*

Applying the necessary adjustments, the results were obtained using a robust linear regression estimated by the ordinary least squares (OLS) method, employing the variables *TARIFF, EXP, MOBILE_HOME,* and *CONCENTRATION*. We have also considered that changes on the interest rates could affect the price of fees charged to consumers, however, when added to the model, this variable was not statistically significant. Besides, we tested the model with quarterly data and for possible seasonality – the regression results of these additional models are available in Appendix C. Except for some variations on the coefficients, the results of all these supplementary regressions did not affect the conclusions evidenced in our main model.

The results of our main regression are presented on Table 11. First, we can see that all variables were statistically significant, since the "t" test, individually, allows us to reject the null hypothesis that the coefficients are statistically equal to zero. In addition, the "F" test, with a value of 76.67165, also allows us to confirm that the independent variables, together, are statistically significant. Furthermore, the model resulted in an adjusted R-square of 0.704983, meaning that the independent variables, together, explain around 70.5% of the *TARIFF* variable, which shows that the model is well adjusted.

As for the coefficients, the positive sign of the variable *EXP* indicates that an increase on the expenses of financial institutions implies in an increase of tariff prices. This variable was equivalent to the marginal costs ($c$) in Wang and Wang's (2018) model and, therefore, we can see that it behaved as predicted in the theoretical model presented in Eq. (1). This is reasonable, once increases in the operational and administrative costs tend to be transferred to consumers in an oligopolistic market such as the Brazilian banking sector.

Regarding *CONCENTRATION,* conceptually speaking this variable represents $n$ in the theoretical model of Eq. (1), which indicated that the number of firms should have a negative effect on prices. However, as explained in Section 3.2, instead of the number of banks we employed a measure of market concentration (HHI), which inverts the expected sign of the coefficient: the higher the concentration is, the higher tariff prices should be, based on the idea that greater market power would lead to an

**Table 11**
Regression results for the adjusted model.

| Variable | Coefficient | Std. Error | t-Statistic | Prob. |
|---|---|---|---|---|
| C | 30.27322 | 3.228822 | 9.375934 | 0.0000 |
| EXP | 0.493498 | 0.131409 | 3.755425 | 0.0003 |
| MOBILE_HOME | -0.729825 | 0.103562 | -7.047248 | 0.0000 |
| CONCENTRATION | -0.010627 | 0.002156 | -4.929107 | 0.0000 |
| **R-squared** | 0.714299 | **Adjusted R-square** | | 0.704983 |
| **F-statistic** | 76.67165 | **Prob(F-statistic)** | | 0.000000 |

*Note*: HAC standard errors & covariance

increase in prices. However, although *CONCENTRATION* was statistically significant, the negative sign shows us that market concentration had a negative impact on tariffs along the analyzed period. This is a surprising result, which could be related to the specific characteristics of markets that have been historically well concentrated.

Finally, the *MOBILE_HOME* variable (our proxy for the invert of transportation costs $t$) returned a negative coefficient, indicating that the expansion on the use of mobile and home banking transactions has led to a reduction on the price of bank fees along the analyzed period. More precisely, the results present us evidence that an increase of one billion in the number of transactions made via mobile or home banking, implies, on average, in a reduction of nearly R$ 0.73 on bank tariff prices. It is interesting to note that the negative coefficient found on the *MOBILE_HOME* variable shows us that the effect of e-banking popularization (reduction in transportation costs) predominated over the effect imposed by increasing expenses, i.e. even though expenses were growing, the pressure generated by the reduction in transportation costs enabled a reduction in tariff prices.

This way we present evidence that the growth in the number of e-banking transactions (reduction of transportation costs) led to a fall in tariffs charged by Brazilian banks. In the past, Veiga and Oliveira (2006) conducted a similar study to this one. However, as mobile and home banking were not widespread technologies at the time, the authors measured digitization through the number of ATMs and electronic only branches. Our results regarding the use of mobile and home banking are similar to their findings, which reinforces that technological progress has the power to increasingly impact the competition conditions of the sector. Therefore, our work advances the empirical literature on the effects of technological progress in finance, showing how digitization can transform a traditional sector in ways that go beyond how customers make their daily banking transactions: it also affects their pockets.

## 5. Conclusions

The Brazilian banking sector has undergone many structural changes in the last two decades, the most of them due to the technological advancements achieved in the market. Investments and expenses in new technology have been growing, a fact that highlights the importance that modernization and the development of new solutions have for the banking sector.

The growth on the supply of online banking services (e-banking) revolutionized the market's dynamic, making the relationship between consumers and financial institutions more agile, automated, efficient and democratic. Besides that, e-banking has the potential to reduce information asymmetry between parties, as well as the transportation costs incurred by consumers on their search for banking services.

In light of that, this work explored, through an econometric study, the impacts that the dissemination of e-banking (mobile and home banking) had on the price of bank tariffs during the period of 2012–2019. Our main argument was that tariff prices should be higher the higher the transportation costs, administrative expenses, and sector concentration were. The number of mobile and home banking transactions was used as the inverse of the transportation cost, since the use of these technologies generates utility and convenience for customers, while decreasing horizontal (locational) differentiation in the banking sector.

The main result we found is that the growth and widespread use of e-banking had a negative impact on the price of fees charged by Brazilian banks. Therefore, in addition to the efficiency and practicality gains provided by the use of mobile and home banking, these technologies also contributed to a reduction in the expenses of customers who pay — usually high — bank fees. This is an interesting result, since it highlights the potential of new technologies to change an oligopolistic and competitive sector, which offers relatively homogenous services and products. The reduction in tariff prices is beneficial to consumers and stimulates competition among the sector's major players, while opening room for new technology-focused financial institutions, such as fintechs.

In addition, there is evidence that banks' expenses are positively related to the fee prices charged by banks. These results confirm our expectations that the price of tariffs should be higher the higher the transportation costs and administrative expenses are. All our results were statistically significant; however, although our expectations were confirmed for the effects of e-banking and expenses, the same was not true for the concentration index (HHI), as we found that higher concentration has not led to higher prices, as one would expect. This relationship should be further analyzed on future research.

It is also important to highlight that our results are based on the existence of cointegration among at least 3 variables of the model. Thus, by saying that the series are cointegrated, we can state that there is a long-term relationship in the data. It is noteworthy, however, that in the short term the variables might behave differently, and such analysis could also be interesting. To analyze short-term results, future research could apply the Error Correction Mechanism (ECM). Finally, there could also be some delay on the impacts of the explanatory variables on bank fees. For instance, financial institutions need a consistent quantity of data to perform their analysis on costumers' digital behavior, which would imply that only after a certain amount of time they would feel comfortable with their decision to lower its fees, as it often takes time for companies to perceive and understand market changes. This could be explored by analyzing the possible existence of lagged effects in the explanatory variables.

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## Declarations of interest

None.

## Appendix A

(see Table A1).

**Table A1**
Data accounted for "Revenue from Services" – code 71700009.

| Code | Account Name |
|---|---|
| 7.1.7.05.00–4 | Revenue from Payment Services |
| 7.1.7.07.00–2 | Revenue from Intermediation on Loans and Financing Between Parties |
| 7.1.7.10.00–6 | Revenue from Administration of Investment Funds |
| 7.1.7.15.00–1 | Revenue from Administration of Funds and Programs |
| 7.1.7.20.00–3 | Lottery Administration Revenue |
| 7.1.7.25.00–8 | Revenue from Administration of Investment Companies |
| 7.1.7.30.00–0 | Revenue from Technical Advisory |
| 7.1.7.35.00–5 | Revenue from Consortium Administration Fees |
| 7.1.7.40.00–7 | Collection Revenue |
| 7.1.7.45.00–2 | Revenue from Commissions For the Placement of Securities |
| 7.1.7.50.00–4 | Revenue from Exchange Brokerage |
| 7.1.7.55.00–9 | Revenue from Administration of Rediscounted Assets |
| 7.1.7.60.00–1 | Revenue from Brokerage of Stock Market Operations |
| 7.1.7.70.00–8 | Revenue from Custody Services |
| 7.1.7.80.00–5 | Revenue from Services Provided to Linked Companies |
| 7.1.7.90.00–2 | Fund Transfer Revenues |
| 7.1.7.94.00–8 | Revenue from Service Packages - Np |
| 7.1.7.95.00–7 | Priority Service Revenues - Np |
| 7.1.7.95.01–4 | Registration Making |
| 7.1.7.95.03–8 | Supply Of 2nd Copy of Magnetic Card with Debit Function |
| 7.1.7.95.04–5 | Supply Of 2nd Copy of Magnetic Savings Account Card |
| 7.1.7.95.05–2 | Exclusion from The Register of Issuers of Checks Without Funds |
| 7.1.7.95.06–9 | Counter-Order, Opposition and Withdrawal of Checks |
| 7.1.7.95.07–6 | Supply of Check Sheets |
| 7.1.7.95.08–3 | Administrative Check |
| 7.1.7.95.10–0 | Certified Check |
| 7.1.7.95.11–7 | Withdrawal from Demand and Savings Account |
| 7.1.7.95.12–4 | Identified Deposit |
| 7.1.7.95.13–1 | Supply of Monthly Statement or Period Statement |
| 7.1.7.95.14–8 | Supply of Microfilm, Microfiche or Similar |
| 7.1.7.95.15–5 | Wire/Electronic Transfer |
| 7.1.7.95.16–2 | Scheduled Wire/Electronic Transfer |
| 7.1.7.95.17–9 | Transfer Between Institution's Own Accounts |
| 7.1.7.95.18–6 | Payment/Money Order |
| 7.1.7.95.19–3 | Granting of Advances to Depositors |
| 7.1.7.95.20–3 | Basic Credit Card - Annuity |
| 7.1.7.95.21–0 | Supply Of 2nd Copy of Credit Card |
| 7.1.7.95.22–7 | Use of Service Channels for Cash Withdrawal - Credit Card |
| 7.1.7.95.23–4 | Payment of Bills Using the Credit Function |
| 7.1.7.95.24–1 | Emergency Credit Assessment - Credit Card |
| 7.1.7.95.25–8 | Manual Exchange Related to International Travel |
| 7.1.7.96.00–6 | Revenue from Differentiated Services – Np |
| 7.1.7.96.01–3 | Investment Funds Management |
| 7.1.7.96.02–0 | Endorsement and Surety |
| 7.1.7.96.03–7 | Valuation, Revaluation and Replacement Of Assets Received As Guarantee |
| 7.1.7.96.04–4 | Exchange |
| 7.1.7.96.05–1 | Differentiated Credit Card - Differentiated Annuity |
| 7.1.7.96.06–8 | Prepaid Card |
| 7.1.7.96.07–5 | Brokerage Involving Securities, Derivatives and Custody |
| 7.1.7.96.99–6 | Other Differentiated Services – Np |
| 7.1.7.97.00–5 | Special Services Revenue – Np |
| 7.1.7.97.10–8 | Overdrafts Rates |
| 7.1.7.97.11–5 | Overdrafts Rates – Mei |
| 7.1.7.97.99–5 | Other Rates |
| 7.1.7.98.00–4 | Bank Fee Revenue – Lp |
| 7.1.7.98.01–1 | Register |
| 7.1.7.98.02–8 | Deposit Accounts |
| 7.1.7.98.03–5 | Resources Transfer |
| 7.1.7.98.04–2 | Credit Operations |
| 7.1.7.98.05–9 | Overdraft Rate - Legal Person |
| 7.1.7.98.99–4 | Other Bank Fee Revenues – Lp |
| 7.1.7.99.00–3 | Revenue from Other Services |

(see Table A2).

**Table A2**
Data accounted for "Administrative Expenses" – code 81700006.

| Code | Account Name |
|---|---|
| 8.1.7.03.00–3 | (-) Water, Energy and Gas Expenses |
| 8.1.7.06.00–0 | (-) Rent Expenses |
| 8.1.7.09.00–7 | (-) Expenses for Renting Goods |
| 8.1.7.12.00–1 | (-) Communication Expenses |
| 8.1.7.15.00–8 | (-) Philanthropic Contribution Expenses |
| 8.1.7.18.00–5 | (-) Honorary Expenses |
| 8.1.7.18.10–8 | (-) Fiscal Council |
| 8.1.7.18.30–4 | (-) Management and Board of Directors |
| 8.1.7.21.00–9 | (-) Maintenance and Conservation Expenses |
| 8.1.7.21.10–2 | Permanent Assets |
| 8.1.7.21.20–5 | Non-Financial Assets Held for Sale - Owned |
| 8.1.7.21.30–8 | Non-Financial Assets Held for Sale - Received |
| 8.1.7.21.40–1 | Leased Goods |
| 8.1.7.21.90–6 | Others |
| 8.1.7.24.00–6 | (-) Material Expenses |
| 8.1.7.27.00–3 | (-) Staff Expenses - Benefits |
| 8.1.7.30.00–7 | (-) Staff Expenses - Social Charges |
| 8.1.7.30.10–0 | (-) Service Guarantee Fund |
| 8.1.7.30.50–2 | (-) Social Security |
| 8.1.7.30.60–5 | (-) Supplementary Social Security |
| 8.1.7.30.99–7 | (-) Others |
| 8.1.7.33.00–4 | (-) Staff Expenses - Provents |
| 8.1.7.36.00–1 | (-) Staff Expenses - Training |
| 8.1.7.37.00–0 | (-) Internship Remuneration Expenses |
| 8.1.7.39.00–8 | (-) Data Processing Expenses |
| 8.1.7.42.00–2 | (-) Expenditure on Promotion and Public Relations |
| 8.1.7.45.00–9 | (-) Advertising and Marketing Expenses |
| 8.1.7.48.00–6 | (-) Expenditure on Publications |
| 8.1.7.51.00–0 | (-) Insurance Expenses |
| 8.1.7.54.00–7 | (-) Financial System Services Expenses |
| 8.1.7.57.00–4 | (-) Third Party Services Expenses |
| 8.1.7.60.00–8 | (-) Surveillance and Security Services Expenses |
| 8.1.7.63.00–5 | (-) Specialized Technical Services Expenses |
| 8.1.7.66.00–2 | (-) Transportation Expenses |
| 8.1.7.69.00–9 | (-) Tax Expenses |
| 8.1.7.72.00–3 | (-) Foreign Travel Expenses |
| 8.1.7.75.00–0 | (-) National Travel Expenses |
| 8.1.7.77.00–8 | (-) Fines Applied by The Central Bank |
| 8.1.7.81.00–1 | (-) Fund Administration Fee Expenses |
| 8.1.7.99.00–0 | (-) Other Administrative Expenses |

(see Table A3).

**Table A3**
Accounts derived from the Credit Guarantee Fund.

| Account Name |
|---|
| Term/Time deposits |
| Bill of exchange |
| Real estate letters |
| Mortgage letters |
| Mortgage bonds |
| Demand deposits |
| Savings deposits |
| Agribusiness letters of credit |
| Non-mov. check deposits |
| Compromised operations |

(see Table A4).

**Table A4**
Descriptive Statistics.

| Variable | Mean | Median | Std Dev. | Min | Max |
|---|---|---|---|---|---|
| **Raw Data** | | | | | |
| REVENUE FROM SERVICES (R$ Billion) | 8.51 | 8.41 | 1.58 | 5.65 | 12.35 |
| ADMINISTRATIVE EXPENSES (R$ Billion) | 14.12 | 14.29 | 1.86 | 10.62 | 18.95 |
| IGP-M | 609.69 | 620.55 | 81.23 | 474.14 | 759.11 |
| NUMBER OF ACCOUNTS (million) | 283.95 | 281.56 | 32.16 | 227.02 | 338.76 |
| **Model Data** | | | | | |
| TARIFF | 20.18 | 19.96 | 1.62 | 17.32 | 23.97 |
| EXP | 9.52 | 9.37 | 0.69 | 8.01 | 11.55 |
| MOBILE_HOME | 2.93 | 2.81 | 1.17 | 1.21 | 4.98 |
| CONCENTRATION | 1190.64 | 1188.39 | 53.97 | 1086.63 | 1359.29 |

Note: statistics over period 2012M01 - 2019M12 (96 observations), except for the "NUMBER OF ACCOUNTS" which is semiannual (16 observations)

## Appendix B

(see Table B1).

**Table B1**
Pairwise Granger Causality Tests.

| Null Hypothesis: | F-Statistic | | | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | LAG 1 | LAG 2 | LAG 3 | LAG 4 | LAG 5 | LAG 6 | LAG 7 | LAG 8 | LAG 9 | LAG 10 | LAG 11 | LAG 12 |
| EXP does not Granger Cause TARIFF | 0.3855 | 2.3758 * | 1.78850 | 2.3460 * | 1.8834 * | 2.600 * * | 1.9224 * | 2.42 * * | 1.9462 * | 1.9256 * | 3.07 * ** | 2.222 * * |
| TARIFF does not Granger Cause EXP | 0.0129 | 1.34134 | 0.54338 | 1.33437 | 1.15947 | 2.227 * * | 1.8811 * | 1.934 * | 2.127 * * | 3.70 * ** | 3.36 * ** | 1.7282 * |
| MOBILE_HOME does not Granger Cause TARIFF | 38 * ** | 15.2 * ** | 7.27 * ** | 5.25 * ** | 2.578 * * | 1.8774 * | 1.58686 | 1.6145 | 2.021 * * | 2.302 * * | 2.430 * * | 2.220 * * |
| TARIFF does not Granger Cause MOBILE_HOME | 26 * ** | 12.9 * ** | 9.02 * ** | 7.35 * ** | 6.23 * ** | 5.66 * ** | 5.52 * ** | 5.1 * ** | 4.49 * ** | 4.90 * ** | 4.57 * ** | 4.10 * ** |
| CONCENTRATION does not Granger Cause TARIFF | 0.0306 | 1.14089 | 1.14109 | 1.03424 | 0.52994 | 1.62364 | 1.31222 | 1.2007 | 1.24560 | 1.08958 | 0.87511 | 0.64763 |
| TARIFF does not Granger Cause CONCENTRATION | 0.5567 | 0.33986 | 0.20889 | 0.12463 | 0.13065 | 1.01451 | 1.36792 | 1.3265 | 1.52725 | 1.6599 * | 1.41875 | 1.05703 |
| MOBILE_HOME does not Granger Cause EXP | 7.3 * ** | 5.91 * ** | 4.15 * ** | 3.484 * * | 2.961 * * | 3.03 * ** | 3.46 * ** | 3.4 * ** | 2.81 * ** | 2.369 * * | 2.374 * * | 2.57 * ** |
| EXP does not Granger Cause MOBILE_HOME | 12 * ** | 7.90 * ** | 5.62 * ** | 4.21 * ** | 4.44 * ** | 4.64 * ** | 4.39 * ** | 3.8 * ** | 3.49 * ** | 3.61 * ** | 3.24 * ** | 3.54 * ** |
| CONCENTRATION does not Granger Cause EXP | 2.0374 | 1.11706 | 1.70776 | 1.48047 | 1.36582 | 4.20 * ** | 3.74 * ** | 4.1 * ** | 3.55 * ** | 3.33 * ** | 3.25 * ** | 1.868 * * |
| EXP does not Granger Cause CONCENTRATION | 0.1603 | 0.13983 | 0.35683 | 0.74284 | 1.22779 | 0.85974 | 1.25593 | 1.1330 | 1.14490 | 1.52322 | 3.03 * ** | 2.317 * * |
| CONCENTRATION does not Granger Cause MOBILE_HOME | 35 * ** | 17.4 * ** | 11.5 * ** | 8.47 * ** | 6.90 * ** | 5.99 * ** | 5.68 * ** | 4.8 * ** | 4.17 * ** | 3.62 * ** | 3.22 * ** | 2.88 * ** |
| MOBILE_HOME does not Granger Cause CONCENTRATION | 6.35 * * | 3.0325 * | 1.38326 | 1.02501 | 0.70160 | 0.76130 | 0.92107 | 0.8862 | 0.72993 | 0.60903 | 0.51810 | 2.439 * * |

Note: * , * *, * ** indicates rejection at 10%, 5% and 1% respectively.

## Appendix C

(see Table C1).

**Table C1**
Regression results for the model with the inclusion of the interest rate.

| Variable | Coefficient | Std. Error | t-Statistic | Prob. |
|---|---|---|---|---|
| C | 29.82082 | 3.0746 | 9.699091 | 0.0000 |
| EXPENSES | 0.515698 | 0.129041 | 3.996390 | 0.0001 |
| HOME_MOBILE | -0.750084 | 0.106819 | -7.022012 | 0.0000 |
| HHI | -0.010217 | 0.00209 | -4.889493 | 0.0000 |
| INTEREST | -0.245063 | 0.391215 | -0.626416 | 0.5326 |
| **R-squared** | 0.715218 | **Adjusted R-square** | | 0.7027 |
| **F-statistic** | 57.13575 | **Prob(F-statistic)** | | 0.000000 |

(see Table C2).

**Table C2**
Regression results for the model considering quarterly data.

| Variable | Coefficient | Std. Error | t-Statistic | Prob. |
|---|---|---|---|---|
| C | 28.61566 | 4.761746 | 6.009488 | 0.0000 |
| EXPENSES_Q | 0.39146 | 0.178129 | 2.197619 | 0.0364 |
| HOME_MOBILE_Q | -0.800327 | 0.125443 | -6.379998 | 0.0000 |
| HHI_Q | -0.008246 | 0.003301 | -2.497814 | 0.0186 |
| **R-squared** | 0.801427 | **Adjusted R-square** | | 0.780151 |
| **F-statistic** | 37.66873 | **Prob(F-statistic)** | | 0.000000 |

(see Table C3.1).

**Table C3.1**
Seasonality tests.

| Series | QS Test | KW Test | Seasonality |
|---|---|---|---|
| TARIFF | 0.00000 | 0.00000 | TRUE |
| EXPENSES | 0.00000 | 0.00001 | TRUE |
| HHI | 0.00000 | 0.00025 | TRUE |
| HOME_MOBILE | 0.00000 | 0.01824 | TRUE |

Note: QS-test < 0.01 or KW-test < 0.002 indicate the time series as seasonal (Webel and Ollech, 2017)

(see Table C3.2).

**Table C3.2**
Regression results for the model considering seasonality treatment (X-13).

| Variable | Coefficient | Std. Error | t-Statistic | Prob. |
|---|---|---|---|---|
| C | 25.73662 | 2.738218 | 9.399040 | 0.0000 |
| EXPENSES_X13 | 0.221081 | 0.090687 | 2.437835 | 0.0167 |
| HOME_MOBILE_X13 | -0.936513 | 0.103321 | -9.064079 | 0.0000 |
| HHI_X13 | -0.004123 | 0.00229 | -1.800229 | 0.0751 |
| **R-squared** | 0.772689 | **Adjusted R-square** | | 0.765276 |
| **F-statistic** | 104.24370 | **Prob(F-statistic)** | | 0.000000 |

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