# The Macroeconomic Determinants of the Pass-Through from the Market Interest Rate to the Bank Lending Rate in Mozambique

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**Abstract**

This paper employs a linear regression with interaction terms, impulse response functions, and analysis of multiplier effects to identify the macroeconomic determinants of the market-to-bank interest rate pass-through in Mozambique. This paper also looks at how these macroeconomic fundamentals affect the interest rate pass-through mechanism. The study finds incomplete market-to-bank interest rate pass-through and shows that it takes approximately five months for the money market rate to be fully transmitted to the bank lending rate. There is evidence indicating the existence of asymmetry in the interest rate transmission mechanism, and the empirical findings also highlight that GDP growth and inflation are the most important macroeconomic variables influencing the degree of the interest rate pass-through in both the short and long run.

**Key Words** – Mozambique, cointegration, interaction terms, asymmetry, interest rate pass-through, money market rate, bank lending rate
JEL – E43, E44, G21

## 1. Introduction

The market-to-bank interest rate pass-through (IRPT) plays an essential role in the monetary policy transmission mechanism. When setting the policy rate, the central bank expects that it will affect the price of money in the money market and therefore the cost of borrowing from the banks for both households and entrepreneurs. As pointed out by Xu and Chen (2012), monetary policy is effective when a change in the policy rate is passed on to the bank lending rates, which in turn influence domestic demand (consumption and investment) and eventually output. Thus, a quicker and more complete pass-through of monetary policy and money market rates to the banking rates strengthens the monetary policy transmission mechanism (Matemilola et al., 2015).

While many empirical studies have focused on the IRPT mechanism from money market rates to banking rates, only a few studies have addressed the macroeconomic factors that drive this market-to-bank IRPT mechanism. Some examples are De Graeve et al. (2004), Kauko (2005), Bogoev (2010), Gigineishvili (2011), Schluter et al. (2012), and Leroy and Lucotte, (2014). Knowing the factors that drive the IRPT is vital for effective policymaking and financial intermediation. In this paper, I use Mozambican quarterly data for the period 2000:1 and 2016:3 to assess both the short and the long-run pass-through mechanisms by allowing for asymmetry in the adjustment processes. The contribution of this study to the empirical literature is the following. First, to the best of my knowledge, this is the first paper that identifies the determinants of IRPT specifically for Mozambique. Other studies may have dealt with this, but only in a context of a panel of African countries including Mozambique. Thus, I can use the results of other studies for comparison with the results that I obtained from my estimations. Second, while the existing empirical literature is limited to listing the factors that influence the process of the transmission mechanism, this paper delves further into the analysis by seeking to determine how these macroeconomic fundamentals affect the IRPT mechanism.

The results in this paper indicate an incomplete and slow market-to-bank IRPT in Mozambique, and there is evidence supporting the existence of asymmetry in the interest rate adjustment. The results also suggest that GDP growth and inflation are the main macroeconomic factors driving the transmission mechanism in both the short and long run, with positive and negative effects on the transmission mechanism, respectively.

This paper is organised as follows. Section 2 describes some relevant issues on the features of the money market and banking sector in Mozambique. Section 3 reviews the relevant literature on IRPT and its determinants. In Section 4, I describe the methodology procedures. Data issues are described in Section 5. In Section 6, I discuss the empirical results. Finally, conclusions are presented in Section 7.

## 2. The Money Market and Banking Sector in Mozambique

### *2.1 The Money Market*

The monetary policy regime conducted by the Bank of Mozambique (BM) is based on monetary targets. Hence, BM uses the money multiplier to monitor the money supply. When doing so, BM presupposes that the functional relationship between the money supply and inflation is predictable and stable. The primary instrument used by BM for policy making is open-market operations. These operations are conducted on the *Mercado Monetário Interbancário*, the Mozambican money market (MM). Additionally, to strengthen the effectiveness of the monetary policy, BM uses moral persuasion, i.e. BM systematically announces its decisions and policy interventions as well as makes an overview of the development of markets to strengthen the confidence of the players in the Mozambican financial system.

The MM is thin and underdeveloped. In fact, as pointed out by Machava and Brännäs (2015), approximately 78% of Mozambicans do not have access to financial institutions services, and credit is only available to an estimated 3% of the population. Moreover, according to KPMG (2016), more than 80% of the services in the financial system are undertaken by commercial banks. Participation of individuals in this market is limited. The MM was formally established in 1997, three years after the liberalisation of the banking interest rates in Mozambique. Nevertheless, a study by USAID (2007) concluded that the bank lending rates in Mozambique are at a level above what the economic indicators can explain, suggesting incompleteness of the interest rate transmission from policy rate or money market rate to banking rates. The high level of bank lending rates might reflect the relative market power that the commercial banks have in Mozambique (USAID, 2007). Moreover, the banking rate response appears to be weaker for downward rather than upward movements in the money market rate (Massarongo, 2013), suggesting the existence of asymmetric interest rate pass-through. The benchmark interest rate for the MM is the Maputo Interbank Offered

Rate (MAIBOR) with maturity between overnight and one year. When setting the *Facilidade Permanente de Cedência* (FPC), which is the policy rate, the BM aims to influence the MAIBOR and, thus, the banking interest rates. However, a study by Hassan (2007) suggests that treasury bills are the actual benchmark for the Mozambican commercial bank's interest rates. According to Hassan (2007), one of the reasons for this is that the income earned from holding treasury bills is tax-free. Besides the treasury bills (with a maturity of 91 days), others instruments traded in MM are swaps, repo, and reverse repo contracts. Despite efforts from BM to reduce FPC (from 22.54% in 2001 to 7.5% in 2015), the lending rate (from 20.5 % in 2001 to 14.75 % in 2015) continues to be high compared to other countries (the world average lending rate in 2015 was 11.4%).

### *2.2 The Banking Sector*

Currently, the Mozambican financial sector comprises 18 commercial banks. According to BM (2015), the banking sector has been recording growth rates of total assets and outstanding profitability, mainly reflecting an increase in the loan portfolio and efficiencies in the process of financial intermediation. Regarding capital structure, foreign investors, predominantly Portuguese and South Africans, own most of the banks, as is the case for the three largest banks. Credit concentration is high not only compared to more developed countries in southern Africa (e.g., Mauritius and Botswana), but also compared to countries at a similar development stage such as Tanzania, Zambia, and Congo. The high degree of market concentration underpins the banking sector growth by granting high pricing power to the leading banks. These market imperfections might be one of the reasons for the observed high lending rate (USAID, 2007).

A survey from KPMG (2016) points out three factors as constraining the efficiency of banking activity in Mozambique. First is the maturity mismatch between the FPC and commercial bank loans. Second is the fact that big and small banks exchange only a relatively low level of liquidity in the money market, which partly reflects banks’ risk aversion to exposing their financial situation to other competing banks in the sector. Third is the source of credit in foreign currencies, which are limited and subject to restrictions imposed by BM. While playing a key role in the banks’ financing, the bank clients’ share of the total banking system’s deposits is very narrow and is also concentrated in a few large institutional clients.

## 3. Literature Review

### *3.1 Market-to-banking Interest Rate Pass-through*

The IRPT measures the monetary policy transmission mechanism. It quantifies the degree (completeness) and the speed with which changes in monetary policy rates are transmitted to banking interest rates. The size or degree of the IRPT refers to the extent to which changes in the money market rates are passed through to the banking rates in the long run. The speed of the IRPT refers to the length of time required for the banking interest rate adjustment to return to equilibrium. In the majority of the empirical studies, the degree and the speed of adjustment are measured by the size (magnitude) of the coefficients for the IRPT mechanism and by the error correction term, respectively. Hence, the larger the magnitude of the coefficients, the more complete and faster the IRPT mechanism. Note that the completeness and speed of the IRPT are two independent measures. While completeness is likely to be affected by macroeconomic fundamentals and structural factors such as the level of development of the financial sector, the speed of adjustment is closely related to banking efficiency (Schluter et al. 2012).

Table 1: Results on the size of the market-to-bank interest rate pass-through (selected studies)

| | *Country* | *Methodology* | *Short run* | *Long run* | *EC term* |
|---|---|---|---|---|---|
| Das (2015) | India | VECM | 0.053<br>(0.30) | 0.68<br>(0.100) | 0.04<br>(0.009) |
| Aziakpono and Wilson (2013) | South Africa | ECM | – | – | 0.45<br>(–) |
| Misati et al. (2011) | Kenya | ECM | 0.03<br>(–) | 0.33<br>(–) | 0.04<br>(–) |
| | Bulgaria | | – | – | 0.12<br>(–) |
| Karagiannis, et al. (2010) | Slovenia | ECM | – | – | 0.11<br>(–) |
| | Greece | | – | – | 0.13<br>(–) |
| Muller-Spahn (2008) | Germany | ECM | 0.79<br>(–) | 0.94<br>(–) | – |
| Gambacorta (2004) | Italy | VECM | 0.88<br>(–) | 1.01<br>(–) | – |
| Lim (2001) | Australia | AVECM | – | 0.71<br>(–) | – |
| Lowe (1995) | Australia | ECM | – | 0.89<br>(0.08) | – |

Notes: *ECM – Error Correction Model; VECM – Vector Error Correction Model; AVECM – Asymmetric Vector Error Correction Model; "– "and "(–)" stand for coefficient not reported and standard errors not reported, respectively; long run – more than one year; short run – less than one year.*

In general, given differences in the level of financial market development, it is found that the IRPT is stronger and faster in developed economies than in developing ones. Similar features are found for the speed of adjustment of bank rates to changes in the money market rates. The high competitiveness of the bank sector in developed countries makes banks in this group more sensitive to changes in the money market rate relative to those in developing economies. Table 1 summarises the findings of some studies in this regard. There exist two main theories explaining the IRPT mechanism – the monetary policy approach (Sander and Kleimer, 2004) and the cost of funding approach (De Bondt, 2005).

*The monetary policy approach*

The monetary policy approach focuses on the transmission of monetary policy impulses into the financial sector. Hence, this approach is interested in the effect that monetary policy has on banking rates, and it concentrates solely on the question of how closely banking rates follow policy rates (Kwapil and Scharler, 2006). Consequently, the monetary policy transmission mechanism takes place through the relationship between the policy rate and the term structure of the interest rate (the yield curve).[1] Three theories explain the term structure of the interest rate – the expectations theory or unbiased expectations theory (Fisher, 1930), the market segmentation theory (Culbertson, 1957), and the liquidity premium theory (Hicks, 1946).

The three theories separately explain, among other things, the following aspects that are relevant for inferences on monetary policy effectiveness. First, interest rates for shorter and longer maturities tend to move together, suggesting that the direction of the reaction of interest rates along the yield curve does not change with maturity. Second, yield curves typically have positive slopes for lower short-term interest rates and negative slopes for higher short-term interest rates. Indeed, the majority of investors prefer to invest in bonds with short maturity (with higher interest rates) rather than in bonds with long maturity (with lower interest rates). As explained by Demiralp and Yilmaz (2012), the policy rate is likely to directly affect shorter rather than longer-term interest. Thus, the effectiveness of monetary policy is expected to decrease with the interest rate maturity.

[1] In this regard, and as pointed out by Machava and Brännäs (2015), the shape of the yield curve provides relevant information about the potential performance of monetary policy (Estrella and Hardouvelis, 1991).

The monetary policy approach states that changes in the policy rate are fully transmitted to the banking rate, i.e. the IRPT is complete. There are two critical factors for the IRPT completeness.[2] First is the substitutability between bank loans and government bonds on the balance sheet of banks and firms. If bank loans and government bonds are not perfect substitutes (companies cannot just access the capital markets but have to rely on the bank loans), the transmission mechanism of the monetary policy impulse is necessarily linked to bank behaviour.[3] Second, in the structure of the financial system, a close bank-firm relationship tends to weaken the policy-determined interest rate (Sander and Kleimeier, 2000).

*The cost of funding approach*

Market-to-retail pass-through can also be explained using the cost of funding approach, which applies a standard marginal cost pricing model to the financial markets approach (De Bondt, 2005). The cost of funding approach explains how changes in the money market affect the banking rates. This approach is based on the costing-margin (markup) theory (Hall and Hitch, 1939) that views a bank as a profit-maximising intermediate agent providing financial products. The underlying theory is that money market rates reflect marginal or opportunity costs of funds because banks rely on them for short-term borrowing. Consequently, there is a positive relationship between market rates and bank lending rates. They also represent the opportunity costs of deposits for households and enterprises given the alternative possibility of investing in the money market or short-term government securities. The pricing of banks' retail products will also include a premium for the maturity and risk transformation involved in their activities (Gigineishvili, 2011).

Unlike the case in the monetary policy approach, in this alternative approach, the bank interest rates do not adjust in full to changes in the cost of funding because banks have market power because they operate within an oligopoly market. Empirical evidence on the oligopoly structure of the banking sector is extensive. Some examples are the studies by Leuvensteijin et al. (2001), Uchino (2014), and FSAP (2009) for the case of Mozambique.

[2] When the proportion of bank loans in the firm's balance is high, the lending channel is an important element in the transmission mechanism of monetary policy.

[3] Recall that from the theoretical point of view, the relationship between the banking rate and the underlying policy/market rates is typically considered in the economic literature using models of oligopolistic bank competition, where banks act as price-setters (under the assumption that banks have some market power). Thus, they have some mark power to set the level of banking rates regardless of changes in monetary policy.

### 3.2 Asymmetric Adjustment to the IRPT

The hypothesis of asymmetry in the adjustment process is another important issue in the empirical literature on the IRPT. This hypothesis states that in imperfectly competitive banking markets one would expect the size of the market-to-bank IRPT to depend on the direction of the changes in the market interest rates. More precisely, banking rates are more sensitive to upward rather than to downward movements in the monetary policy and money market rate. Likewise, the reaction of banking rates differs depending on whether bank interest rates are below or above equilibrium levels as determined by cointegration relationships (De Bondt, 2005). In the literature, this phenomenon is called interest rate stickiness. Asymmetric adjustments can arise because of asymmetric information (adverse selection and moral hazard). In response to this market imperfection, bank lending rates are likely to become slow to increase and quick to decrease (Popiel, 2015). The policy implication is that, under a scenario of downward stickiness in interest rates an expansionary monetary policy is likely to be more efficient than a tight monetary policy.

Some empirical studies show evidence that asymmetries exist in the IRPT. For example, Lim (2001), Gambacorta (2004), and Das (2015) present evidence of asymmetric adjustments of both deposit and lending rates in Australia, Italy, and India. However, other studies (see, e.g. Karagiannis et al. (2010) and Apergis and Cooray (2015)) have found no evidence for asymmetric adjustment.

The majority of African empirical studies on the asymmetric IRPT are based on panel data. For example, Sander and Kleimeier (2006) found evidence that the speed of adjustment of the lending rate to changes in the market rate in South Africa, Lesotho, Namibia, and Swaziland is asymmetric and that the IRPT takes on average six months to be completed. In a paper studying the link between bank concentration and IRPT in four sub-Saharan countries (Zambia, Nigeria, South Africa, and Namibia), Mangwengwende et al. (2011) also found evidence of asymmetry. In a more recent study estimating a time series model for Nigeria, Fadiran (2012) found evidence in favour of an asymmetric IRPT from discount rate to lending rate.

### 3.3 Determinants of the Market-to-banking IRPT

Since Cottarelli and Kourelis' (1994) pioneering work on the determinants of the IRPT, researchers have been concerned with identifying the structural and macroeconomic factors affecting the monetary policy performance to explain the observed heterogeneity of the IRPT in various countries and markets. Moreover, as noted in the introduction, economists agree that monetary policy effectiveness weakens with interest rate maturities. This weakness might be explained by the fact that longer-maturity interest rates are subject to influence from structural and macroeconomic variables. The empirical literature on the determinants of the IRPT identifies some variables influencing the size and the speed of adjustment of the bank lending rate to changes in the policy rate and the money market rate. For example, in a cross-sectional study covering 70 countries from all regions of the world, including low-income and developed countries, Gigineishvili (2011) identified per capita GDP, inflation, and market volatility as the principal macroeconomic determinants of IRPT. The main results in a selection of studies are presented in Table 2. To my knowledge, there is no empirical research on the determinants of market-to-bank IRPT for an individual African economy. However, there are some studies on the determinants of the lending rate in this group of economies. These studies found similar results to those in studies for advanced economies, i.e., the macroeconomic factors that drive the lending rate are primarily inflation and GDP growth (see, e.g. Malede, 2014 and Mbao et al., 2014).

Table 2: Selected empirical literature on the macroeconomic determinants of the IRPT

| | *GDP* | *Exchange Rate* | *Inflation* | *Market Volatility* | *Financial Market Development* |
|---|---|---|---|---|---|
| Cottarelli and Kourelis (1994) | | | + | + | |
| Mojon (2000) | | | + | + | |
| Sander and Kleimer (2004) | + | | + | + | |
| Gigineishvili (2011) | + | | + | − | |
| Saborowski and Weber (2013) | | + | | | + |

Notes: *The signs "+" and "−" stand for the positive and negative impact of the underlying macroeconomic variable on the market-to-bank interest rate pass-through.*

## 4. The Model

Because the monetary policy approach focuses solely on the question of how closely retail rates follow policy rates and includes no other explanatory variables (Kwapil and Scharler, 2006 ), the model presented in this paper is based on the cost of funding approach. Thus, I propose a simple model based on the interest rate channel of monetary policy transmission where the bank lending rate will be tied the relevant proxy for the measurement of the IRPT. The baseline long-run model is specified in Section 4.1. Section 4.2 introduces an error-correction (EC) model[4] with the bank lending rate also being explained by other macroeconomic variables in the context of both symmetric and asymmetric relationships.

### *4.1.The Standard Pass-through Regression*

This paper documents the macroeconomic determinants of the responsiveness of the bank lending rate ("lending rate" hereafter) to the money market rate ("market rate" hereafter) in Mozambique. As a starting point, consider the following long-run equilibrium relationship between the market rate and the lending rate:

$$r_t = \alpha_0 + \alpha_1 m_t + \mu_t \tag{1}$$

where $r_t$ is the lending rate at time $t$, $m_t$ is the market rate at time $t$, and $\mu_t$ is an error term. $\alpha_0$ is the intercept term and $\alpha_1$ is the coefficient of the long-run market-to-bank IRPT. The pass-through parameter $\alpha_1$ plays an important role in determining the efficiency of transmission, and the size of $\alpha_1$ determines the degree of IRPT. If $\alpha_1 = 0$, there is no IRPT mechanism; if $0<\alpha_1<1$, the IRPT is incomplete, meaning that the demand for banking services is inelastic relative to the lending rate; and if $\alpha_1 = 1$, the IRPT is complete, meaning that the banking and the market rates move precisely in the same direction and proportionally. Such a situation is unlikely to occur in the context of the cost of funding approach because it would require full information and perfect competition in the banking sector.

As mentioned previously, equation (1) describes the long-run equilibrium relationship between market and lending rates. However, short-run equilibrium deviations are likely to arise because banks do not respond to every daily change in the market rates because to do so would be too costly. Moreover, the response of the lending rates to changes in the money market rates might not be instantaneous, and Castro and Santos (2010) found that lags are involved in the market-to-bank IRPT. Indeed, the effects of changes in market rates on lending rates are expected to last over the subsequent periods as the change prevails, i.e. the effects will be "distributed" over periods $t$, $t+1$, $t+2$ … $t+n$. Hence, based on these insights and the result of the Granger causality test (Appendix B), equation (1) is rewritten to account for adjustment lags on the market-to-bank IRPT and thus becomes a single-equation

[4] The model for the EC mechanism might involve more than one equation. The decision on the number of equations consists of performing a Granger causality test (Granger, 1969) to assess whether the causality relationship between the variables is unilateral or bilateral. The results of the Granger causality test (Appendix B) indicate that the transmission is unidirectional and originating from the market rate to the lending rate. Thus, the model in this paper is one of singular EC.

error correction model (ECM):

$$\Delta r_t = \alpha_0 + \sum_{j=0}^{o} \delta_j \Delta r_{t-j} + \sum_{i=0}^{n} \beta_{1+i} \Delta m_{t-i} + \theta \mu_{t-1} + \varepsilon_t ; \tag{2}$$

where $\mu_{t-1} = r_{t-1} - \alpha_1 m_{t-1} - \alpha_0$, $\sum_{j=0}^{o} \delta_j$ is the lagged autoregressive effect of $\Delta r_t$, $\sum_{i=0}^{n} \beta_{1+i}$ measures the market-to-bank short-run lagged IRPT, $o$ and $n$ stand for the lags on lending and market rate changes, respectively, and $\Delta$ denotes the difference operator. This ECM describes how a change in the lending rate in period $t$ responds to a change in market rate in period $t$ plus a correction for the discrepancy between the level of the lending rate and its equilibrium value at time $t-1$. $\mu_{t-1}$ is the error correction term, which is the lagged disturbance term from the regression in levels (equation (1)). During periods of disequilibrium, $\mu_{t-1}$ is a non-zero term and its value corresponds to how far the lending rate is from the equilibrium during the time $t$. Thus, an estimate of $\theta$ provides information on the speed of adjustment back to equilibrium (Harris et al., 1995). If the ECM is appropriate, then $-1 < \theta < 0$. A $\theta$ value close to $-1$ implies a quick correction, while a small value close to zero suggests a slow adjustment toward the long-run steady state.

Now, and based on the discussion in Section 3.2, equations (1) and (2) can be transformed to account for asymmetries. To this end, I follow Shin et al. (2014) to specify the following asymmetric long-run equation of market-to-bank IRPT:

$$r_t = \alpha_0 + \alpha_1^+ m_t^+ + \alpha_1^- m_t^- + \nu_t \tag{3}$$

where $m_t^+ = \sum_{k=1}^{t} \Delta m_k^+$ and $m_t^- = \sum_{k=1}^{t} \Delta m_k^-$ are the partial cumulative sums of positive and negative changes in the market rates (measured by $\alpha_1^+$ and $\alpha_1^-$), respectively, and where $\Delta m_t^+ = \max(\Delta m_t, 0)$ and $\Delta m_t^- = \min(\Delta m_t, 0)$. Both $\alpha_1^+$ and $\alpha_1^-$ are expected to be positive as stated in the cost of funding approach. Hence, symmetric IRPT is just a particular case of the asymmetric IRPT when $\alpha_1^+ = \alpha_1^-$. If $\alpha_1^+ < \alpha_1^-$, there is upward interest rate stickiness, meaning that the lending rate is less elastic to upward than to downward movements in the market rates. If $\alpha_1^+ > \alpha_1^-$, lending rate is more elastic to upwards than to downwards movements in the market rate, i.e. there is a downward stickiness in the interest rate. This rigidity in the interest rate is the source of the asymmetry adjustment in the IRPT. The $\nu_t$ is a white noise process and $c$ is the intercept. Following Shin et al. (2014), the corresponding standard ECM with asymmetric coefficients for both short and long-run dynamics is of the following form:

$$\Delta r_t = \alpha_0 + \sum_{j=0}^{o} \delta_j \Delta r_{t-j} + \sum_{i=0}^{n} \beta_{1+i}^{+} \left( \Delta m_{t-i}^{+} \right) + \sum_{i=0}^{n} \beta_{1+k}^{-} \left( \Delta m_{t-i}^{-} \right) + \theta \mu_{t-1} + \varepsilon_t \qquad (4)$$

where $\mu_{t-1} = r_{t-1} - \alpha_1^{+} m_{t-1}^{+} - \alpha_1^{-} m_{t-1}^{-} - \alpha_0$. Equations (1)–(4) address the first objective of this paper, i.e. the question of whether there is a market-to-bank IRPT in both the short and the long run in Mozambique. However, these equations might not do well in explaining the short and long-run IRPT because they do not consider the effects of fundamental macroeconomic factors. As mentioned in Section 3.3, some economic factors, including those measuring the financial system soundness and bank-level factors, might be important in explaining short-run changes in the lending rate, even though they might have no long-run effects. However, limitations in the data impede an empirical assessment of the impacts of the financial and bank-level factors on the market-to-bank IRPT in Mozambique. The follow up is to address the fundamental macroeconomic factors driving the market-to-bank IRPT.

### *4.2. Assessing the Determinants of the Market-to-bank IRPT*

This paper employs a time series analysis to assess the macroeconomic factors driving the market-to-bank IRPT. According to the literature, the standard empirical approach to determinants of the IRPT mechanism uses a two-step estimation procedure within a panel or cross-section data framework (Ozdemir and Altinoz, 2012). In the first stage, both short and long-run IRPT coefficients are estimated on individual time series, and then, in the second stage, the estimated IRPT coefficients are related to variables that represent the macroeconomic, financial, and bank sector features of each country's IRPT.[5]

While it is straightforward in a panel or cross-sectional setting to use the two-stage procedure to estimate the effects of macroeconomic factors on the IRPT, an alternative approach is required for the single time series context. In a conventional linear time-series framework, the pass-through coefficient is constant across time periods, yielding only one observation. To overcome this limitation, I use a linear regression model with interaction terms, which is interpretable as a time-varying parameter model. This procedure enables a

[5] Due to data unavailability, the analysis of the determinants of the market-to-bank IRPT will be restricted to macroeconomic factors. Moreover, the usage of individual bank interest rate data would be more convenient; however, institutional restrictions in the Mozambican banking sector make this type of information unavailable in a systematic manner.

one-step estimation approach that captures some aspects of the two-step approach adopted for the panel and cross-sectional data, which are left out in a regular linear regression. Estimators obtained from this type of model are comparable to those based on the panel or cross-section data sets. In both cases, they represent the average effects with the only difference being that inferences based on time series are aggregated impacts for a particular country, while conclusions in the panel setting are made for a group of economies or to disaggregated data from a specific country.

The inclusion of interaction terms results in a linear model with more than one slope. Therefore, the interpretation of the results from the model is more complicated than in the simple linear model (Jaccard and Turrissi, 2003). However, one advantage of having a model with interaction terms is the flexibility for making inferences. In other words, in a model with interaction terms, one can assess the effects of control variables on the relationship between the two variables of primary interest, and one can analyse how the size of the macroeconomic effect unfolds for small and large levels of money market rates. To capture these features, consider equation (5), which is an extension of equation (1), where the slope (the IRPT coefficient) is dependent on the variables described above. Note that in the model specification, the macroeconomic variables are assumed to affect both the lending rate and the IRPT coefficient.

$$r_t = \alpha_0 + \alpha_{1t} m_t + \sum_{j=1}^{n} \varphi_j z_{jt} + \varepsilon_t \tag{5}$$

where $\alpha_{1t} = \alpha_1 + \sum_{j=1}^{n} \phi_j z_{jt}$. In equation (5), $z_{jt}$ stands for the selected macroeconomic variables, $j = 1, 2 \ldots n$ is the index for the individual macroeconomic factors, and $\varepsilon_t$ is a white noise residual. By rewriting equation (5), a model with interaction terms is obtained:

$$r_t = \alpha_0 + \alpha_1 m_t + \sum_{j=1}^{o} \phi_j \left( z_{it} m_t \right) + \sum_{i=1}^{n} \varphi_j z_{it} + \varepsilon_t \tag{6}$$

Equation (6) has three sets of parameters where $\alpha_1$ is the coefficient for the long-run market-to-bank IRPT without the influence of the macroeconomic variables, $\phi_j$ is the coefficient for the interaction between the money market rate and the selected macroeconomic variable $j$

and measures the direct impact of each macroeconomic determinant on the IRPT coefficient, and $\varphi_j$ is the impact of the macroeconomic variables on the bank lending rate[6].

The ECM model associated with the long-run specification in equation (6) can be written as follows:

$$\Delta r_t = \alpha_0 + \sum_{j=0}^{o} \delta_j \Delta r_{t-j} + \sum_{i=0}^{p} \beta_{1+i} \Delta m_{t-i} + \sum_{j=1}^{n}\sum_{i=0}^{q} \gamma_{ij} \Delta\left(z_{t-i} m_{t-i}\right) + \sum_{j=1}^{n}\sum_{i=0}^{s} \pi_{ij} \Delta z_{t-i} + \theta \mu_{t-1} + \zeta_t$$

$$\text{where } \mu_{t-1} = r_{t-1} - \alpha_0 - \alpha_t m_{t-1} - \sum_{i=1}^{n} \phi_i \left(z_{it-1} m_{t-1}\right) - \sum_{i=1}^{n} \varphi_i z_{it-1} \tag{7}$$

where $p$, $q$, and $s$ are the number of lags for short-run changes in the money market rate, for the interaction terms, and for the control variables, respectively. The value of $\theta$ is assumed to be constant. Recall that, as discussed in Section 3.1, the value of $\theta$ is influenced by banking sector factors rather than macroeconomic variables. In line with the procedure in the baseline model, we can also specify the corresponding long-run asymmetric model, and an EC representation with asymmetric coefficients for short-run dynamics is as follows:

$$r_t = \alpha_0 + \alpha_1^+ m_t^+ + \alpha_1^- m_t^- + \sum_{j=1}^{n} \phi_j^+ \left(z_{jt} m_t\right)^+ + \sum_{j=1}^{n} \phi_j^- \left(z_{it} m_t\right)^- + \sum_{j=1}^{n} \varphi_j z_{jt} + \mu_t \tag{8}$$

$$\begin{aligned} \Delta r_t = {} & \alpha_0 + \sum_{j=0}^{o} \delta_j \Delta r_{t-j} + \sum_{i=0}^{p} \beta_{1+i}^+ \Delta m_{t-i}^+ + \sum_{i=0}^{p} \beta_{1-i}^- \Delta m_{t-i}^- + \\ & + \sum_{i=1}^{n}\sum_{j=0}^{q} \gamma_{ij}^+ \Delta\left(z_{t-i} m_{t-i}\right)^+ + \sum_{i=1}^{n}\sum_{j=0}^{q} \gamma_{ij}^- \Delta\left(z_{t-i} m_{t-i}\right)^- + \sum_{i=1}^{n}\sum_{j=0}^{s} \pi_{ij} \Delta z_{t-i} + \theta \mu_{t-1} + \zeta_t \end{aligned} \tag{9}$$

where

$\mu_{t-1} = r_{t-1} - \alpha_0 - \alpha_1^+ m_{t-1} - \alpha_1^- m_{t-1} - \sum_{j=1}^{n} \phi_j^+ \left(z_{jt-1} m_{t-1}\right)^+ - \sum_{j=1}^{n} \phi_j^- \left(z_{jt-1} m_{t-1}\right)^- - \sum_{j=1}^{n} \varphi_i z_{jt-1}$, and where $\sum_{i=0}^{p} \beta_{1-i}^+$ and $\sum_{i=0}^{p} \beta_{1-i}^-$ measure the short-run positive and negative effects, respectively, of

[6] One might argue that there is a good chance of a high correlation between $m_t$ and $z_{it} m_t$ in equation (6), which possibly increases multicollinearity, thereby increasing the size of the standard errors and making it less likely that the coefficient on the interaction term will be significant (Brambor et al., 2006). Nevertheless, Allison (2012) has pointed out that the p-values of $z_t m_t$ are not affected by multicollinearity, meaning that the estimated standard errors are efficient. As explained in Williams and Richard (2016), the multicollinearity in models with interaction terms can be removed by centering $m_t$ and $z_{it}$, subtracting their means, before creating the product $z_{it} m_t$. They argue that the p-value for $z_{it} m_t$ will be the same regardless of whether the variables are centred or not and that all of the results for the other estimates (including the $R^2$) will be the same in either case. Recall that within the context of interaction terms, the coefficients in the models no longer indicate the average effect of a variable as they do in a standard linear regression (Brambor et al., 2006). Thus, when estimating models with interaction terms, researchers are interested in marginal effects. The marginal effects for equation (6) are derived in Appendix C.

changes in $m_t$ on $r_t$ without the influence of the macroeconomic variables. $\sum_{j=0}^{q}\gamma_{ij}^{+}$ and $\sum_{j=0}^{p}\gamma_{ij}^{-}$ stand for the positive and negative effects, respectively, of the macroeconomic variables on the short-run IRPT. Similarly, $\sum_{j=0}^{s}\pi_{ij}$ measures the short-run effect of the macroeconomic variables on the interest rate. $\theta$ is the parameter for the speed of adjustment of the lending rate toward its equilibrium in the long run.

### *4.3. Estimation Procedures*

Because macroeconomic variables are likely to be non-stationary variables, the estimators obtained from equations (1)–(8) are expected to be spurious (Verbeek, 2012). Thus, the first step of the estimation procedure is to identify the order of integration of each variable to identify the stationarity proprieties of the underlying variable. The second step consists of testing for cointegration among the variables to determine the long-run relationship between them. The conventional procedure is Johansen's cointegration method (Habibi and Rahim, 2009). However, I follow the bounds testing approach (see Appendix D) proposed by Pesaran and Shin (1995) and Pesaran et al. (2001) for three reasons. First, there is a need to account for the small sample size problem. Second, given that equations (6) and (8) involve macroeconomic variables (where some might be stationary), the use of the bounds testing approach is suitable because it is applicable irrespective of whether the underlying regressors are purely first-order integrated, I(1), strictly zero-order integrated, I(0), or a mixture of both. Third, structural breaks might impact the cointegration process, and as will be discussed in the next section, there is a suspicion of some structural breaks during the period covered in this study. The bounds test procedure allows for the inclusion of these types of dummy variables.

## 5. Data

The quarterly data for the period 2001:1–2016:3 are from the BM. The lending rate is the average of the loan rates charged to the public (households and entrepreneurs) by the leading commercial banks in Mozambique. The overnight market rate and the three-month treasury bills rate represent the average cost of borrowing money between banks and between banks and the central bank, respectively.[7] The source of the data for the selected macroeconomic indicators used in this study (GDP growth and inflation) is the website of the *Instituto Nacional de Estatística* (the Mozambican bureau of statistics).

Figure 1: Bank lending rate and market rate in Mozambique, 2001:1-2016:3

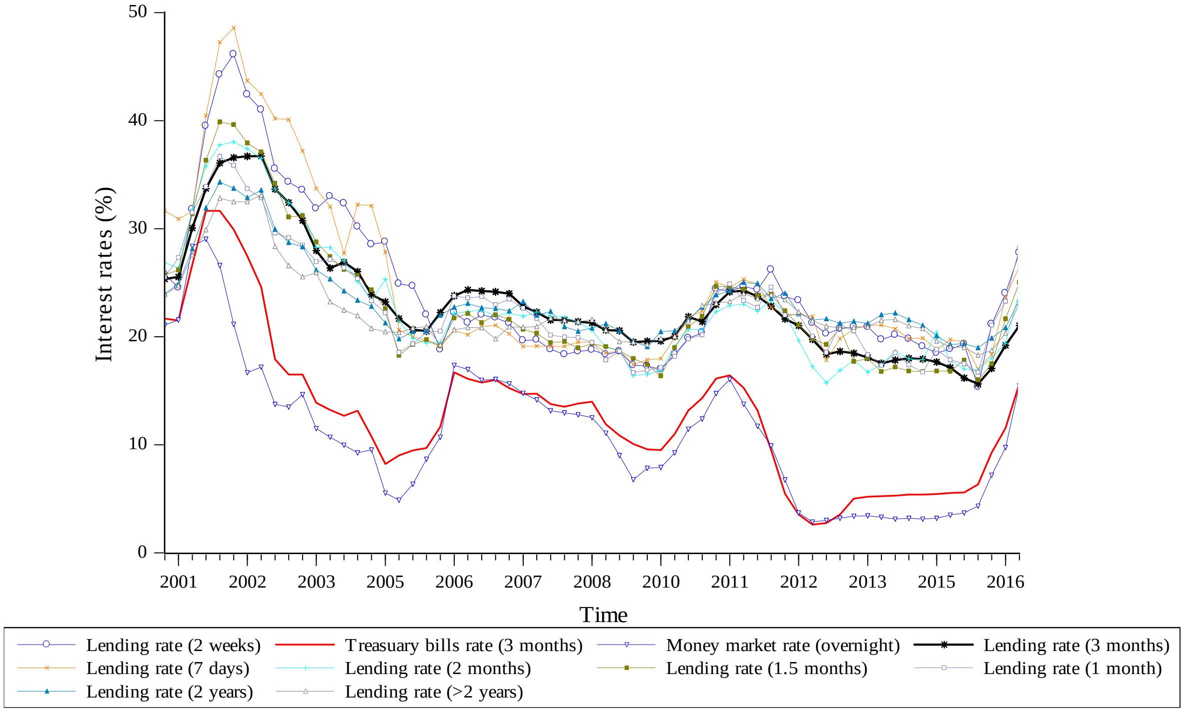



Table A provides summary statistics of relevant interest rates and selected macroeconomic variables. The standard deviations (5.24 for the bank lending rate and 6.78 for the treasury bills rate) are relatively high. As stated in the empirical literature, this is an indicator that banks typically do not adequately adjust the retail rates when the market rates change (Kwapil and Scharler, 2006). Figure 1 plots the average quarterly interest rates for different maturities charged by the commercial banks in Mozambique for loans along with the average overnight market rate and treasury bill rate for transactions in the money market.

Figure 1 shows that the maturities of the instruments play an important role in determining the level of the bank lending rate. The lower the maturity, the higher the level of the bank lending rate, which is noteworthy because as shown in Machava and Brännäs (2015) the yield curve in Mozambique is positively sloped. Figure 1 provides insights into how the selected interest rates unfold over time. All interest rates display a downward trend, which seem to move closer together over the entire sample. However, three possible structural breaks can be

[7] Repo and reverse repo are not included due to data unavailability. Repo and reverse repo contracts are not traded frequently.

identified as a result of changes in monetary policy. The first potential breakpoint was November 2005 (2005Q4) when BM introduced a new methodology for setting the monetary policy rate. This involved indexing the policy rate to the average rates of the treasury bill auctions. As a result, the policy rate increased from 11.5% in October 2005 to 19% in March 2006. The second potential breakpoint is August 2011 (2011Q3) when, after a period of a tight monetary policy, the central bank adopted a looser monetary policy that aimed to induce banks to charge lower lending rates. The last potential breakpoint is June 2015 (2015Q3) when, after four years of loose monetary policy, the central bank increased the policy rate as a response to inflation pressures caused by the sharp appreciation of the US dollar against the domestic currency (the metical).

As shown in Figure 1, there are two candidates for use as a proxy for the cost of funding – the overnight money market rate and the treasury bills rate. One should identify which of the two accurately measures the marginal cost price, and to this end, I follow the standard procedure in the literature that involves comparing the correlations between the lending rates and the market rates across maturities and lags. Additionally, the selected lending rate and the proxy of the cost of funding must have a comparable maturity to avoid distortions resulting from maturity mismatches (De Bondt, 2005). Table C in Appendix C presents the results of the correlations computed across maturities and lags over the period 2001:1 to 2016:3. The correlation coefficients in terms of level vary between 0.70 and 0.83 for the overnight money market rate and between 0.80 and 0.93 for the treasury bills rate, implying that the latter is the best proxy for the cost of funding for commercial banks in Mozambique. The lags with higher correlations vary between one and three quarters, suggesting that bank lending rates do not collectively react at the same speed to changes in the treasury bills rate. In first differences, the correlation coefficients are lower with a range of 0.44 and 0.67 for the overnight money market rate and a range of 0.46 and 0.68 for the treasury bills rate. Similar to Hassan (2007), with a few exceptions, the bank lending rates react immediately to changes in the money market rates. The conclusion is that treasury bills rates are the most appropriate marginal cost prices, and thus this is the interest rate to be used for modelling the market-to-bank IRPT because of its accurate reflection of the marginal funding costs faced by commercial banks.

## 6. Empirical Results

The first step of a time series analysis is to perform unit root tests. Thus, the Augmented Dickey-Fuller (ADF) unit root test is performed for each variable to assess its order of integration. The null hypothesis is that the series is nonstationary, i.e., it has a unit root. First, the series is tested in levels, and if the null hypothesis is not rejected the series is differenced to make it stationary. Otherwise, the model can be estimated with the series in levels. The results of the ADF unit root test are shown in Table 3 and indicate that the null hypothesis of non-stationarity for the levels of the variables is not rejected for the lending rate and treasury bills rate. The GDP is stationary for every test specifications, while inflation is stationary in the case where only the intercept is accounted for. In first differences, the null hypothesis is rejected for all variables at the 5% significance level. Thus all variables are found to be stationary, and the Autoregressive Distributed Lagged–Error Correction Model (ARDL-ECM) methodology can be performed because none of the series is I(2).

Table 3: ADF unit root test

| Variables | ADF specification | t-statistics (Level) | t-statistics (First Difference) |
|---|---|---|---|
| Lending rate | No intercept | −0.104 | −5.842 *** |
| | Intercept only | −1.167 | −5.803*** |
| | Intercept and Trend | −0.955 | −5.780 *** |
| Treasury bills rate | No intercept | −0.876 | −3.843** |
| | Intercept only | −2.444 | −3.813** |
| | Intercept and Trend | −3.267* | −3.775** |
| GDP growth | No intercept | −1.840* | −3.358** |
| | Intercept only | −6.339*** | −7.521*** |
| | Intercept and Trend | −7.633*** | −7.465*** |
| Inflation rate | No intercept | −0.518 | −5.122*** |
| | Intercept only | −4.228** | −5.085** |
| | Intercept and Trend | −2.687 | −5.137** |

Note: *rejection of the null hypothesis of unit root at the 1% level (***), 5% level (**), and 10% level (*)*

Now follow the cointegration tests (bounds test). Before proceeding to cointegration testing per se, it is convenient to test for potential breakpoints. The Chow breakpoint test (Chow, 1960) is performed to test whether the known three points (2005Q4, 2011Q3, and 2015Q3) are points of structural breaks or not. In the Chow test, the null hypothesis of no structural change is tested against the alternative hypothesis of structural change. The results, reported in Table F (Appendix F), indicate that 2005Q5 is the only breakpoint. Hence, the bounds test, with a dummy variable for the breakpoint, is performed for each of the ARDL model specifications in equations (E1)–(E4) in Appendix E. The results are reported in Table 4.

Table 4: Results of the bounds test

| | Dependent variable: Lending rate | | |
|---|---|---|---|
| | F-statistic | I(0) Bound | I(1) Bound |
| Equation A1 | 6.40** | 4.94 | 5.73 |
| Equation A2 | 5.81** | 3.79 | 4.85 |
| Equation A3 | 10.80*** | 4.29 | 5.61 |
| Equation A4 | 5.64*** | 3.23 | 4.35 |

Note: *the null hypothesis is that no long-run relationship exists; rejection of the null hypothesis at *** $p < 0.01$ and ** $p < 0.05$.*

The results indicate evidence for cointegration between the lending rate and the treasury bills rate for all four model specifications considered in this paper. In fact, as shown in Table 4, for the first case (equation A1), the F-statistic for the bounds test is 6.40, and this exceeds the 5% critical value for the upper bound. Accordingly, the null hypothesis of "no long-run linear relationship" is rejected. Thus, the lending rate and the treasury bills rate are cointegrated, i.e. there is evidence that these two series move together in the long run. Similar results are found for the models estimated by equations (A2)–(A4). Note that one of the requirements for the accuracy of the coefficients estimated under the ARDL framework is model stability. Cumulative sum tests (Turner, 2010) indicate that all four model specifications used for the bounds test are stable, meaning that the estimators measure the relationship between the cointegrated variables accurately (Appendix G).

*The IRPT*

The next step is to estimate the long-run IRPT in equations (1), (3), (6), and (8) and the short-run IRPT in equations (2), (4), (6), and (8). The results are reported in Table 5. After estimating each model, the Wald test (Wald, 1939) is performed to test for asymmetry ($\alpha_1^+ \neq \alpha_1^-$ in the long run, and $\beta_1^+ \neq \beta_1^-$ in the short run) and for the pass-through completeness ($\alpha_1 = 1$, $\alpha_1^+ = 1$, $\alpha_1^- = 1$ in the long run, and $\beta_1 = 1$, $\beta_1^+ = 1$, $\beta_1^- = 1$ in the short run). All models are tested against serial correlation using the Durbin–Watson test (Durbin and Watson, 1951). Results for equation (1) indicate that there is a long-run equilibrium relationship between the treasury bills rate and the lending rate, i.e. a 10 percentage point increase in treasury bills rate will result in a long-run increase of 7.1 percentage points in the lending rate. However, this finding may be questioned for two reasons.

Table 5: Long-run interest rate pass-through

(Dependent variable: bank lending rate, $r_t$ )

| | Eq. (1) | Eq. (3) | Eq. (6) | Eq. (8) |
|---|---|---|---|---|
| $m_t$ | 0.71*** | | 1.47** | |
| | (0.176) | | (0.406) | |
| $m_t^+$ | | 0.86*** | | 2.29*** |
| | | (0.141) | | (0.411) |
| $m_t^-$ | | 0.55*** | | 1.40*** |
| | | (0.077) | | (0.215) |
| $m_t \times \inf_t$ | | | −0.05** | |
| | | | (0.015) | |
| $m_t^+ \times \inf_t$ | | | | −0.042*** |
| | | | | (0.009) |
| $m_t^- \times \inf_t$ | | | | −0.040*** |
| | | | | (0.008) |
| $m_t \times gdp_t$ | | | −0.20* | |
| | | | (0.104) | |
| $m_t^+ \times gdp_t$ | | | | −0.35*** |
| | | | | (0.087) |
| $m_t^- \times gdp_t$ | | | | −0.24*** |
| | | | | (0.059) |
| $\inf_t$ | | | 1.07** | 0.65*** |
| | | | (0.300) | (0.135) |
| $gdp_t$ | | | 4.98 | 1.49 |
| | | | (1.653) | (0.991) |
| Constant | 10.20*** | 14.21*** | −7.23 | 0.96 |
| | (2.2849) | (1.868) | (7.487) | (3.375) |
| $\Delta(b\_2005q4)$ | | −6.30*** | −3.835** | −6.61*** |
| | | (1.164) | (1.470) | (0.990) |
| R-squared | 48.6 | 58.1 | 64.3 | 81.8 |
| Observations | 64 | 63 | 64 | 64 |
| DW-statistic | 1.74 | 1.87 | 2.14 | 2.11 |
| Wald Test for Long-Run Models | | | | |
| $\alpha_1^+ = \alpha_1^-$ | | 14.84*** | | 6.39*** |
| | | (0.000) | | (0.000) |
| $\alpha_1^+ = 1$ | | 1.01 | | 9.78** |
| | | (0.315) | | (0.002) |
| $\alpha_1^- = 1$ | | 34.51*** | | 3.38* |
| | | (0.000) | | (0.066) |
| $\alpha_1 = 1$ | 2.65 | | 1.329 | |
| | (0.103) | | (0.249) | |

Notes: *Newey–West standard errors for the model specifications;*
*P-values in parentheses for the Wald test (null hypothesis: no asymmetry);*
*Coefficients statistically significant at *** p < 0.01, ** p < 0.05, and * p < 0.1*

Table 6: Short-run interest rate pass-through

(Dependent variable: change in bank lending rate, $\Delta r_t$ )

| | Eq. (2) | Eq. (4) | Eq. (7) | Eq. (9) |
|---|---|---|---|---|
| $\Delta r_{t-1}$ | | | | −0.24** (0.082) |
| $\Delta m_{t-1}$ | 0.32*** (0.057) | | 0.41*** (0.076) | |
| $\Delta m_t^+$ | | 0.29*** (0.036) | | 0.91*** (0.138) |
| $\Delta m_t^-$ | | 0.19*** (0.022) | | 0.56*** (0.086) |
| $\Delta(\inf_t \times m_t)$ | | | −0.013*** (0.003) | |
| $\Delta(\inf_t \times m_t)^+$ | | | | −0.017*** (0.003) |
| $\Delta(\inf_t \times m_t)^-$ | | | | −0.016*** (0.003) |
| $\Delta(gdp_t \times m_t)$ | | | 0.06** (0.025) | |
| $\Delta(gdp_t \times m_t)^+$ | | | | −0.14 (0.034) |
| $\Delta(gdp_t \times m_t)^-$ | | | | −0.10 (0.023) |
| $\Delta \inf_t$ | | | 0.30*** (0.059) | 0.19** (0.054) |
| $\Delta gdp_t$ | | | −0.24 (0.571) | 0.60 (0.410) |
| $ec_{t-1}$ | −0.14** (0.056) | −0.34*** (0.056) | −0.28*** (0.055) | −0.40*** (0.050) |
| Constant | 1.38* (0.806) | 4.88** (1.303) | −2.023 (1.805) | 0.39 (1.453) |
| $\Delta(b\_2005q4)$ | | −2.16*** (0.468) | −1.07** (0.484) | −2.66*** (0.495) |
| Observations | 64 | 63 | 64 | 62 |
| DW-statistic | 1.74 | 1.87 | 2.14 | 2.29 |
| Wald test for short-run models | | | | |
| $\beta_1^+ = \beta_1^-$ | | 22.81*** (0.000) | | 43.08*** (0.000) |
| $\beta_1^+ = 1$ | | 377.29*** (0.000) | | 231.74*** (0.000) |
| $\beta_1^- = 1$ | | 1315.5*** (0.000) | | 0.350 (0.554) |
| $\beta_1 = 1$ | 140.07*** (0.000) | | 59.37*** (0.000) | |

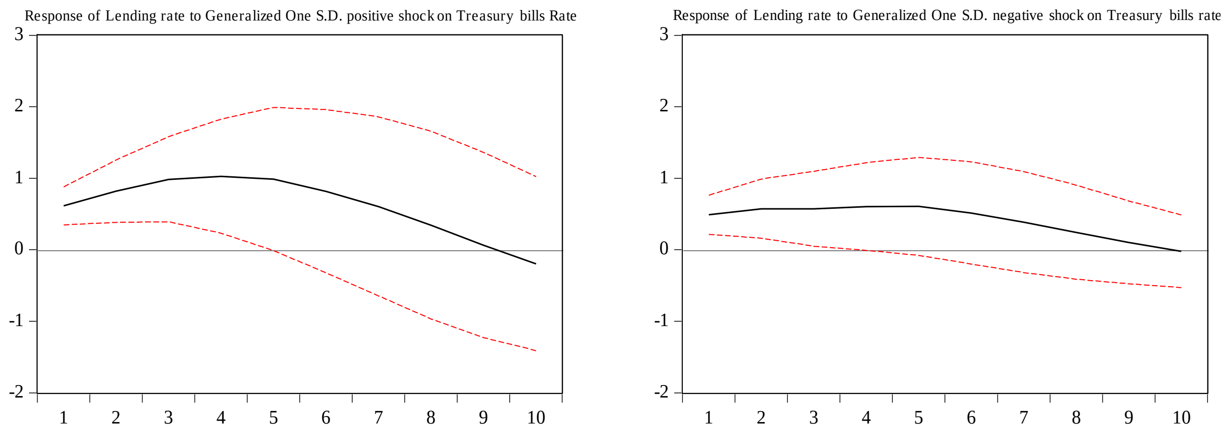
Notes: *Newey–West standard errors for the model specifications;*
*P-values in parentheses for the Wald test (null hypothesis: no asymmetry);*
*Coefficients statistically significant at *** p < 0.01, ** p < 0.05, and * p < 0.1*

First, recall the bounds test result, which suggests the existence of nonlinear long-run co-movement between the variables of primary interest in this study and that the regression in equation (1) might suffer from a misspecification problem because the effects of macroeconomic factors are neglected. Hence, moving to equation (3), which accounts for asymmetry effects, the results show that the impact of the treasury bills rate's cumulative upward movement (0.86) is larger than the corresponding cumulative downward movements (0.55). In other words, in the long run, the lending rate is more responsive to positive rather than negative shocks in the treasury bills rate. When the baseline model is extended by adding macroeconomic factors (equation (6)), the long-run IRPT increases to 1.47. Finally, the results for the model incorporating asymmetric adjustment and the effects of macroeconomic variables (equation (8)) support the finding of downward rigidity in equation (3). The null hypothesis of non-symmetry in the Wald test is rejected in both equations (3) and (8), and these results are in line with the findings by Lim (2001), Karagiannis et al. (2010), and Apergis and Cooray (2015). The Wald test is also performed to check whether the coefficients of IRPT are complete ($\alpha_1 = 1$). The results suggest complete IRPT in equations (1) and (6). However, when accounting for asymmetry (equations (3) and (8)), the IRPT is incomplete. The Durbin–Watson (DW) statistics are high, mainly for the models with interaction terms, and thus serial correlation, and therefore multicollinearity, is not an issue in these models.

The results for the ECM are presented in Table 6. Here, equation (2) indicates a small short-run market-to-bank IRPT of 0.32. When allowing for asymmetric adjustments as in equation (4), the short-run IRPT coefficients for the short-run upward and downward adjustments are 0.29 and 0.19, respectively, suggesting a downward rigidity. As in the long-run model specification, the addition of macroeconomic factors to the simple short-run model (equation (7)) increases the coefficient of the short-run IRPT. The results of the Wald test reject the null hypothesis of non-asymmetry. However, there is evidence in favour of pass-through incompleteness in all model specifications.

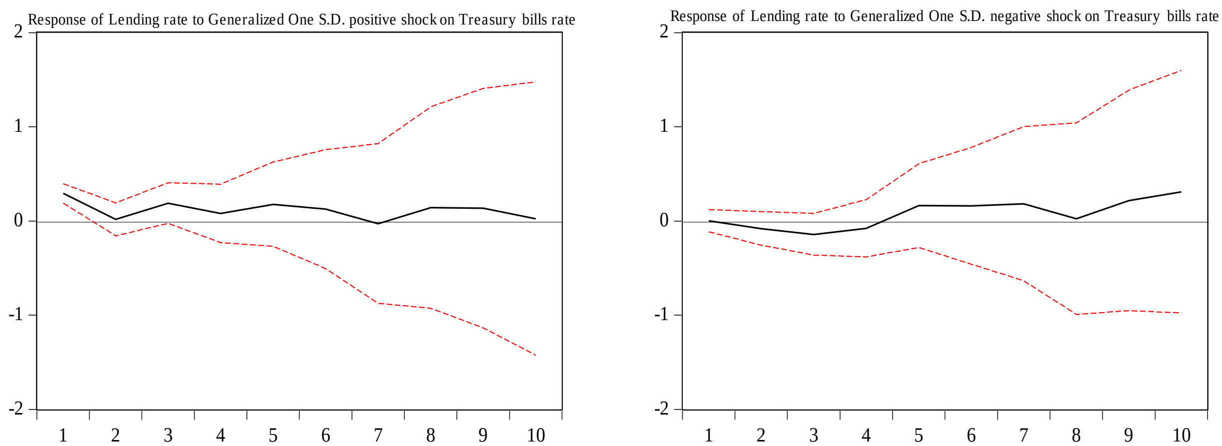
Based on the discussion in Section 4 regarding the estimation of linear regressions with interaction terms, it is incorrect to interpret the coefficients in equations (3) and (8) as the effects of the explanatory variable on the dependent variable. An alternative way to make inferences on these asymmetry dynamics is to plot the impulse response functions (IRFs) of the lending rate against changes in the treasury bills rate.

Figure 2: Impulse response functions for the baseline model (Equation (5))




Note: *The upper and lower curves are 95% confidence intervals, and the black (solid) line is the estimated IRF. The macroeconomic variables used for the estimation are GDP growth and Inflation. SD stands for standard deviation.*

Figure 3: Impulse response functions for the model with macroeconomic variables (Equation (9))




Note: *The upper and lower curves are 95% confidence intervals, and the black (solid) line is the estimated IRF. The macroeconomic variables used for the estimation are GDP growth and Inflation. SD stands for standard deviation.*

Figures 2 and 3 show the IRF of the lending rate for both positive and negative shocks in the treasury bills rate. The IRFs in Figures 2 and 3 show that the lending rate is more responsive to positive shocks to the treasury bills rate rather than negative shocks. Therefore, the results of asymmetry adjustment found in equations (4) and (9) are confirmed. These results are in line with Das (2015) and Apergis and Cooray (2015) who found that the lending rate adjusts more quickly to monetary tightening than to loosening in India and the US and the UK and Australia, respectively. In the African context, these results are consistent with the findings of Fadiran (2012) for Nigeria and of Mangwengwende et al. (2011) for selected sub-Saharan countries (South Africa, Nigeria, Namibia, and Zambia).

*The macroeconomic determinants of the IRPT and the long-run multiplier effect*[8]

After confirming the existence of the IRPT mechanism, the next step is identifying the macroeconomic factors driving the IRPT. As discussed in Section 4.2, the coefficients estimated in Tables 5 and 6 fail to measure the total long-run IRPT, and the multiplier effect framework is required to address this issue. In this context, the multiplier effect refers to the increase in the lending rate that follows from a 1% increase in the treasury bills rate, conditional on the impact of a given macroeconomic factor. To better understand the nature of these marginal effects, it is often instructive to draw line graphs of the estimated marginal effects. The figures of the marginal effects plot the calculated values of the impact of $m_t$ on $r_t$ in both short and long-run cases and conditional on each of the selected macroeconomic variables, against the sample values of $m_t$ in ascending order. As suggested by equation (9), the estimated multiplier effect of $m_t$ on $r_t$ can be evaluated at any chosen values of $m_t$ and $z_{it}$. Computing the multiplier effects requires the selection of specific fixed values of the variables $m_t$ and $z_{it}$. Standard choices for these values, as in this paper, are the median (50$^{\text{th}}$ percentile), the first quartile (25$^{\text{th}}$ percentile), and the third quartile (75$^{\text{th}}$ percentile) (Royston, 2013). The results are displayed in Figures 4 and 5. Starting from Figure 6, the multiplier effect of the treasury bills rate on the lending rate conditional on inflation is positive in the sense that the higher the level of inflation, the larger the coefficient of the long-run market-to-bank IRPT. This result is in line with the findings of Sander and Kleimer (2004) and Gigineishvili (2011). However, this effect becomes narrower for higher levels of the treasury bills rate. The marginal effect of GDP growth on the market-to-bank IRPT is opposite to that of inflation, i.e., GDP growth weakens the IRPT mechanism, and its negative impact becomes stronger for higher levels of the treasury bills rate. Again, this result is in line with the findings of Gigineishvili (2011) and Schluter et al. (2012). The same type of analysis is done for the short-run relationships (Figures 6 and 7). The results are similar to those in the long run, but now both inflation and GDP growth changes have an increasing effect on the market-to-bank IRPT mechanism.

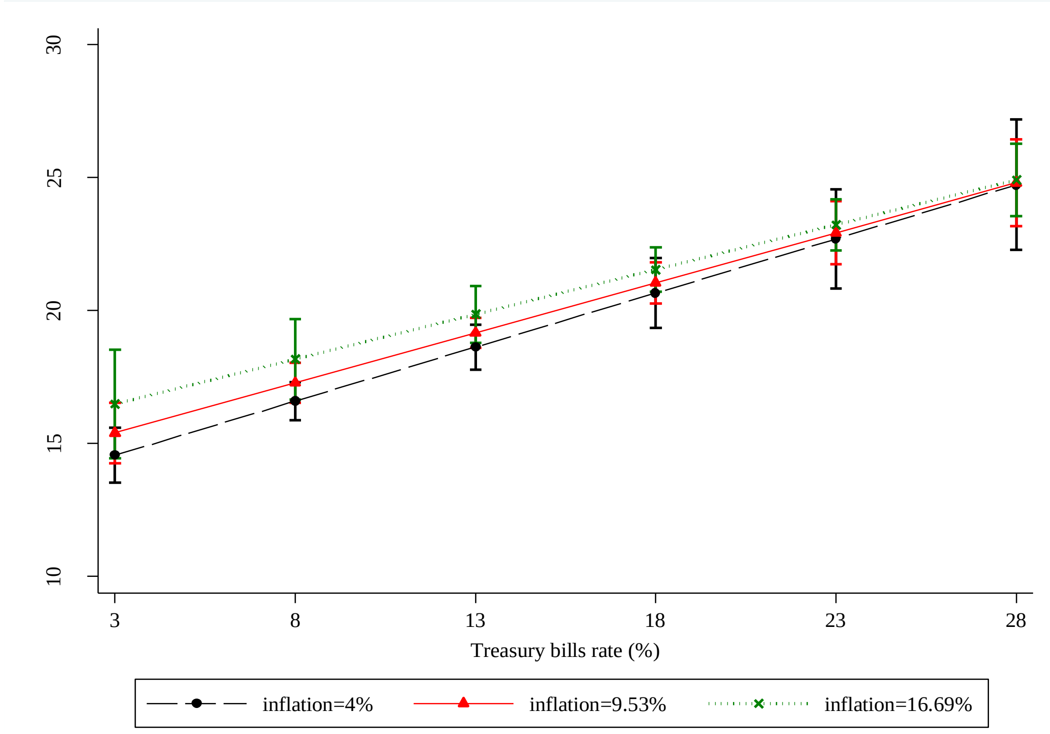
[8] The assessment of the macroeconomic variables playing a role in the determination of the market-to-bank IRPT in Mozambique started by relating variables identified in the literature to both the banking rates and market rates. The following procedure was to retain in the model those variables found to be statistically significant in explaining the IRPT. The following variables were selected based on data availability: exchange rate volatility, GDP growth, inflation, and interest rate volatility.

*The speed of adjustment*

Another relevant characteristic of the IRPT concerns the length of time the adjustment process takes when both variables move toward the long-run equilibrium. The results in Table 6 indicate that for the model in equation (2) the speed of adjustment is slow (0.14), meaning that only 14% of any disequilibrium between the lending rate and treasury bills rate is corrected within a quarter. When considering an extension of the baseline model to account for asymmetries in the short-run dynamics, the speed of adjustment increases to 0.34 (equation (4)). Similar features are observed when the macroeconomic factors are added to the baseline model, and the speed of adjustment increases to 0.28 (equation (7)). Finally, equation (9) indicates that after a movement in the treasury bills rate, the lending rate will increase 40% per quarter until equilibrium is restored. This adjustment process toward the long-run equilibrium takes approximately five months.[9] Although this speed of adjustment is slow in absolute values, it is relatively faster than the average speed of six months for South Africa, Lesotho, Namibia, and Swaziland as found by Sander and Kleimeier (2006). Note that, as required, the coefficient for the EC is negative and significant in all model specifications. The speed of adjustment increases significantly from equation (3) to equation (9), probably reflecting the relevant role of both asymmetry dynamics in interest rate adjustment and the importance of the macroeconomic fundamentals on the IRPT process.

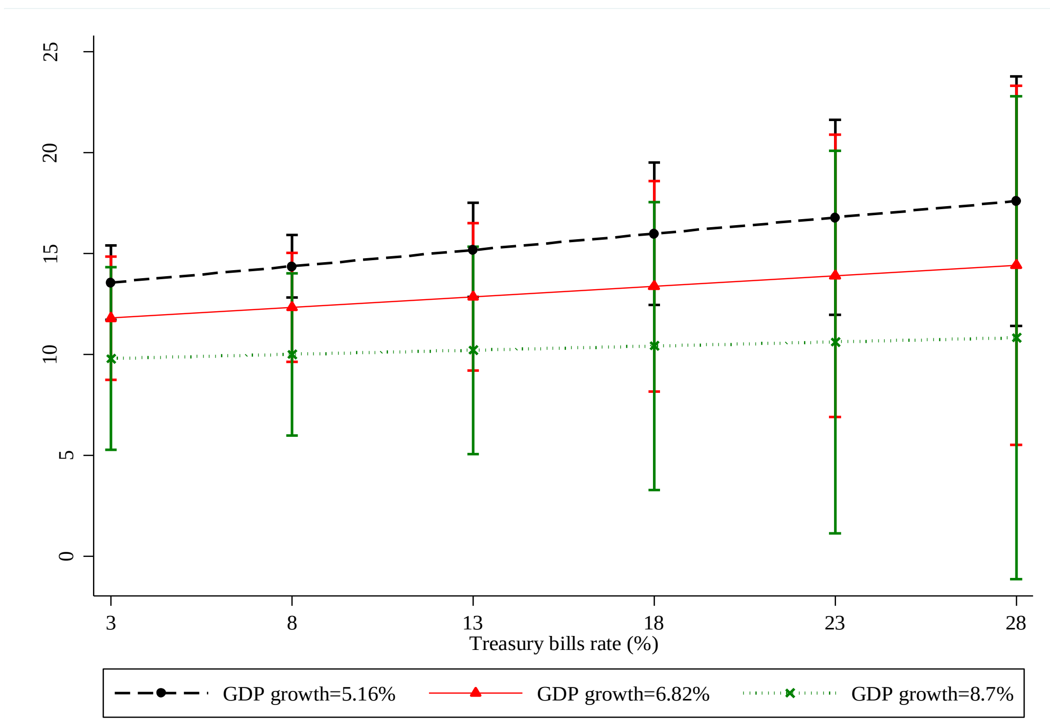
[9] This result is obtained based on the formula for the half-life measure (Fanelli and Paruoloy, 2007). The formula for the half-life measure is $\frac{\ln 2}{\theta}$ , and given the coefficient of adjustment obtained from the ECM (0.29), the result is 1.73 periods.

Figure 4: Long-run multiplier effect of the treasury bills rate on the lending rate conditional on inflation



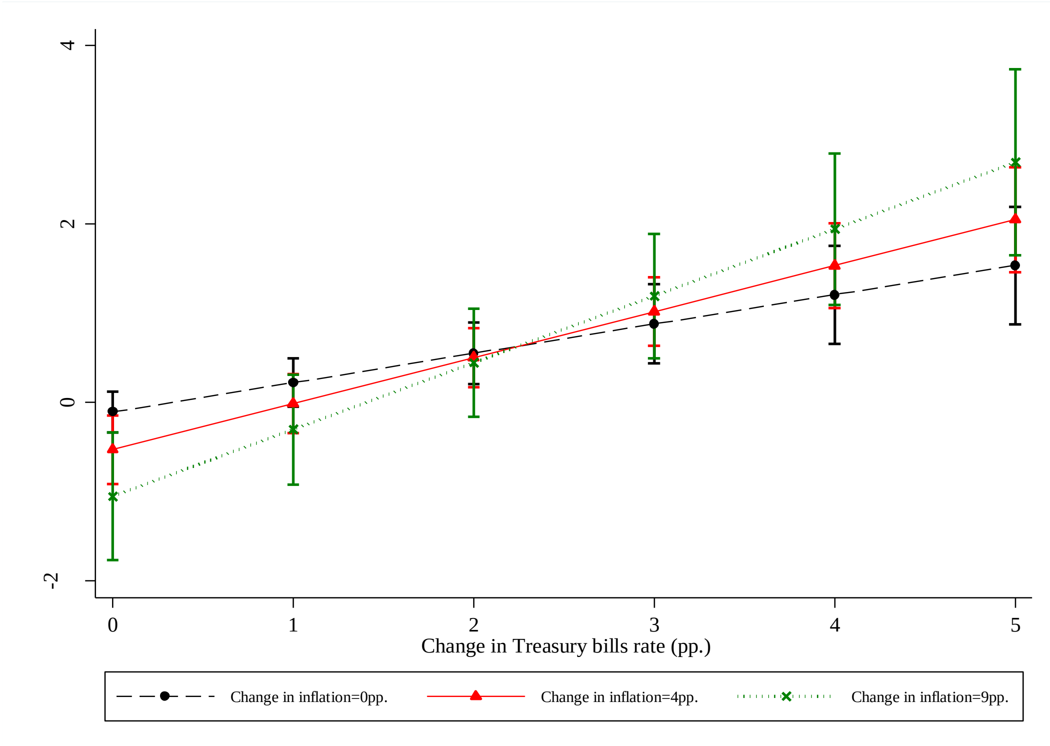

Note: Predictive margins with 95% confidence intervals

Figure 5: Long-run multiplier effect of the treasury bills rate on the lending rate conditional on GDP growth



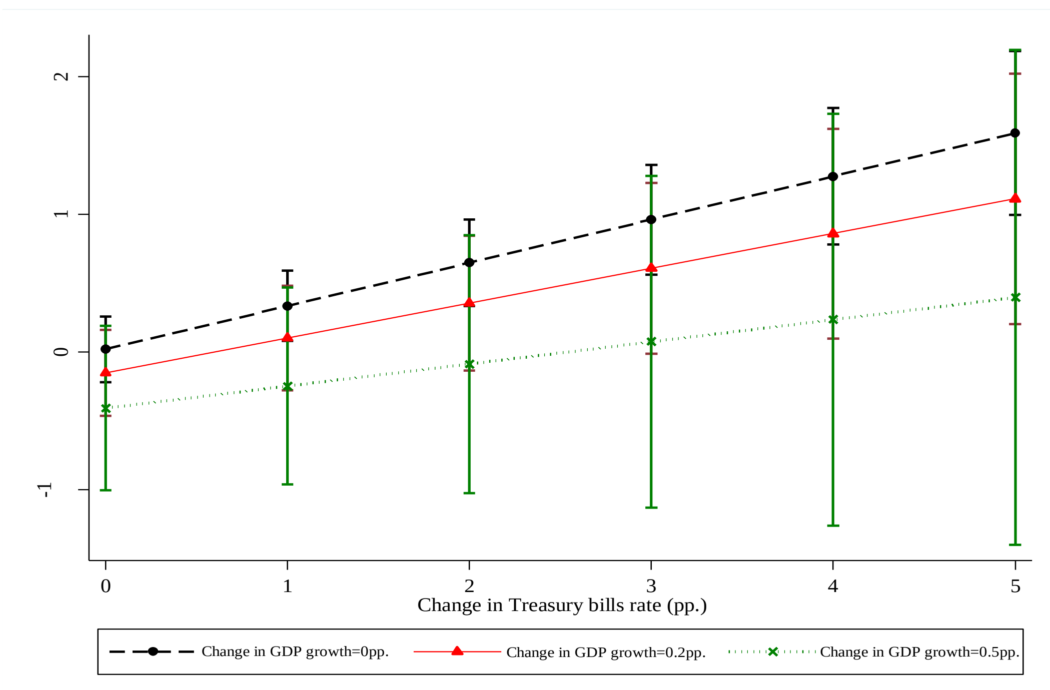

Note: Predictive margins with 95% confidence intervals

Figure 6: Short-run multiplier effect of the treasury bills rate on the lending rate conditional on inflation




Note: Predictive margins with 95% confidence intervals;

Figure 7: Short-run multiplier effect of the treasury bills rate on the lending rate conditional on GDP growth




Note: Predictive margins with 95% confidence intervals;

## 7. Conclusion

This paper employed a regular regression with interaction terms to assess the primary macroeconomic factors driving the market-to-bank IRPT in Mozambique. Cointegration and error correction techniques with both linear and nonlinear specifications were used to study the short and long-run equilibrium dynamics between the lending rate and the money market rate.

The market-to-bank IRPT in Mozambique is found to be incomplete, which is in line with previous panel setting studies, implying that the market-to-bank transmission of interest rates is not efficient. Incomplete pass-through might be the result of a lack of competition in the banking sector. Although the bank lending rate seems to be more responsive to upward than to downward changes in the money market rate, there is not enough evidence to support either the existence of asymmetric interest rate adjustment or the hypothesis of downward rigidity. Thus, no inference regarding efficiency can be made for differences between expansionary and tight monetary policies. The results also indicate that the bank lending rate might deviate from its equilibrium path for extended periods of time. On average, the response of the bank lending rate to changes in the money market rate takes approximately five months.

The empirical findings also highlight that inflation and GDP growth are important for the degree of the IRPT in the transmission of treasury bills rates to lending rates in both short and long-run time periods. Thus, the finding in the standard empirical literature based on panel data receives support from the alternative approach used in this study. In fact, for the Mozambican case, inflation strengthens, and GDP growth weakens the interest rate transmission mechanism. Exchange rate and interest rate volatility do not seem to play a role in the IRPT in Mozambique. Instead, asymmetry in interest rate adjustment and macroeconomic variables seem to play an essential role in the speed of IRPT.

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**Appendix A: Descriptive Statistics**

Table A: Descriptive statistics of the bank lending and money market rates, 2000:1 to 2016:3

| | Lending rate ($r_t$) | Treasury bills rate ($m_t$) | GDP growth ($gdp_t$) | Inflation rate ($\inf_t$) |
|---|---|---|---|---|
| Mean | 22.93 | 13.04 | 7.23 | 9.37 |
| Standard Deviation | 5.24 | 6.78 | 3.29 | 5.61 |
| Minimum | 15.64 | 2.60 | 0.45 | 1.54 |
| Maximum | 36.71 | 31.60 | 16.3 | 23.67 |
| 25$^{th}$ percentile | 19.41 | 9.01 | 5.16 | 4.00 |
| 50$^{th}$ percentile | 21.55 | 13.14 | 6.82 | 9.53 |
| 75$^{th}$ percentile | 24.24 | 16.10 | 8.70 | 19.69 |

## Appendix B: Test for Causality Relationship

The $r_t$ ( $m_t$ ) is said to Granger-cause $m_t$ ( $r_t$ ) if $m_t$ ( $r_t$ ) can be better predicted using the historical values of both $r_t$ ( $m_t$ ) and $m_t$ ( $r_t$ ) than it can by using the history of $m_t$ ( $r_t$ ) alone. The Granger causality test can be performed by estimating the following Vector Autoregressive (VAR) model:

$$\begin{aligned} r_t &= a_0 + a_1 r_{t-1} + \ldots + a_p r_{t-p} + b_1 m_{t-1} + \ldots + b_p m_{t-p} + \mu_t \\ m_t &= c_0 + c_1 m_{t-1} + \ldots + c_p m_{t-p} + d_1 r_{t-1} + \ldots + d_p r_{t-p} + v_t \end{aligned} \tag{A.1}$$

where $a_o$ and $c_o$ are the intercepts, $v_t$ is the disturbance term, and $p$ is the number of lags. Then, testing $H_0 : b_1 = b_2 = \ldots b_p = 0$ against $H_1 : b_1 \neq b_2 \neq \ldots b_p \neq 0$ is a test that $m_t$ does not Granger-cause $r_t$. Similarly, testing $H_0 : d_1 = d_2 = \ldots d_p = 0$ against $H_0 : d_1 \neq d_2 \neq \ldots d_p \neq 0$ is a test that $r_t$ does not Granger-cause $m_t$. In both cases, rejection of the null hypothesis implies that there is Granger causality. Hence, if the analysis indicates the existence of bidirectional causality, then a multivariate model (a system of equations) is required to describe the equilibrium dynamics. Otherwise, a single-equation error-correction model is suitable for explaining the short and long-run dynamics between variables. The results are presented in Table B.

Table B: Granger causality Wald tests

| *Dependent Variable* | *Diagnostics* | *Explanatory Variable* | |
|---|---|---|---|
| | | Treasury bills rate | Lending rate |
| Lending rate | $chi^2$ | 13.40 | |
| | df | 2 | |
| | Prob > $chi^2$ | 0.001** | |
| Treasury bills rate | $chi^2$ | | 3.09 |
| | df | | 2 |
| | Prob > $chi^2$ | | 0.213 |

*Note: ** Null hypothesis of no Granger causality rejected at 5% statistical significance*
*df – degrees of freedom; Prob - probability*

## Appendix C: Correlation Analysis

Table C: Correlation analysis between the bank lending rate and the treasury bills rate

| Bank lending rate | Money market rate (overnight) | | Treasury bills rate (3 months) | |
|---|---|---|---|---|
| | Correlation | Lag (quarters) | Correlation | Lag (quarters) |
| *Variables in level* | | | | |
| 7 days | 0.70 | 3 | 0.80 | 2 |
| 15 days | 0.70 | 3 | 0.84 | 2 |
| 1 month | 0.78 | 2 | 0.87 | 1 |
| 1.5 month | 0.80 | 2 | 0.91 | 2 |
| 2 months | 0.83 | 2 | 0.93 | 2 |
| 3 months | 0.83 | 2 | 0.93 | 2 |
| 6 months | 0.80 | 2 | 0.87 | 2 |
| 1 year | 0.82 | 2 | 0.90 | 2 |
| 2 years | 0.80 | 3 | 0.86 | 1 |
| > 2 years | 0.77 | 2 | 0.83 | 2 |
| *Variables in first difference* | | | | |
| 7 days | 0.44 | 0 | 0.46 | 0 |
| 15 days | 0.46 | 0 | 0.55 | 0 |
| 1 month | 0.67 | 0 | 0.64 | 0 |
| 1.5 month | 0.63 | 0 | 0.63 | 1 |
| 2 months | 0.49 | 0 | 0.53 | 0 |
| 3 months | 0.55 | 0 | 0.68 | 0 |
| 6 months | 0.44 | 0 | 0.46 | 2 |
| 1 year | 0.54 | 0 | 0.62 | 1 |
| 2 years | 0.62 | 0 | 0.63 | 0 |
| > 2 years | 0.44 | 0 | 0.56 | 1 |

In Table C, the dependent variable (the bank lending rate) is related to the two candidates to be used as a proxy for the money market rate (the overnight money market rate and the treasury bills rate) over different maturities (from 7 days to more than 2 years). The optimal lag length selected for every computed correlation is based on the Schwartz Bayesian information criteria. The results indicate that compared to the overnight money market rates, the treasury bills rates are, on average, more correlated to the bank lending rate in terms of both level and first difference. Thus the treasury bills rate is the best proxy for measuring the money market rate.

**Appendix D: Multiplier Effects on the Long-Run Pass-Through**

Here the marginal effects (the multiplier effects) for equation (8) are estimated. The marginal effect of an individual explanatory variable is a function that gives the change in the mean of the dependent variable in response to a one-unit increase in the value of the explanatory variable, holding the effects of all other explanatory variables constant (Williams, 2016). Following Norton et al. (2004), the marginal effects (the dynamic multiplier effects) are obtained from the partial derivatives. Hence, the marginal effect of a change in $m_t$ on $r_t$ is:

$$\frac{\partial r_t}{\partial m_t} = \frac{\partial r_t\left(r_t \middle| m_t, z_{it}\right)}{\partial m_t} = \beta_1 + \sum_{i=1}^{n} \phi_i z_{it} \quad \text{(D.1)}$$

In other words, while the simple model in equation (1) states that $m_t$ has a constant ($\beta_1$) effect on $r_t$, the model with interaction terms indicates that the effect of a change in $r_t$ depends on $z_{it}$. Thus, the standard error of interest is:

$$\sigma_{\frac{\partial r_t}{\partial m_t}} = \sqrt{\mathrm{var}\left(\hat{\beta}_1\right) + \mathrm{z}_{1t}^2\,\mathrm{var}\left(\hat{\phi}_1\right) + \mathrm{z}_{2t}^2\,\mathrm{var}\left(\hat{\phi}_2\right) + \ldots + \mathrm{z}_{nt}^2\,\mathrm{var}\left(\hat{\phi}_n\right) + 2\sum\nolimits_1^n z\,\mathrm{cov}\left(\hat{\beta}_1 \hat{\phi}_i\right)} \quad \text{(D.2)}$$

Analogously, the effect of the change on $z_{it}$ on $r_t$ depends on $m_t$ as shown below:

$$\frac{\partial r_t}{\partial z_{it}} = \frac{\partial r_t\left(r_t \middle| z_{it}, m_t\right)}{\partial z_{it}} = \phi_i m_t + \varphi_i \quad \text{(D.3)}$$

The impact of a marginal change in $z_t$ on the marginal effect of $m_t$, i.e. the interaction effect $\frac{\partial r_t}{\partial\left(m_t z_{it}\right)}$, is then obtained by taking the derivative of (C.1) with respect to $z_{it}$:

$$\frac{\partial r_t}{\partial m_t z_{it}} = \frac{\partial r_t\left(r_t \middle| m_t z_{it}\right)}{\partial m_t z_{it}} = \phi_i \quad \text{(D.4)}$$

**Appendix E: Bounds Test**

The bounds test is based on the autoregressive distributed lagged (ARDL) framework, thus concerning equations (1), (3), (6), and (8) and the following corresponding ARDL models:

$$\Delta r_t = \beta_0 + \sum_{j=1}^{q} \psi_j \Delta r_{t-j} + \sum_{i=0}^{p} \beta_1 \Delta m_{t-i} + \eta_0 r_{t-1} + \eta_1 m_{t-1} + \varepsilon_t \tag{E.1}$$

$$\Delta r_t = \beta_0 + \sum_{j=1}^{q} \psi_j \Delta r_{t-j} + \sum_{i=0}^{p} \beta_{1+i}^{+} \left(\Delta m_{t-i}^{+}\right) + \sum_{i=0}^{p} \beta_{1+i}^{-} \left(\Delta m_{t-i}^{-}\right) + \eta_0 r_{t-1} + \eta_1^{+} m_{t-1}^{+} + \eta_1^{-} m_{t-1}^{-} + \varepsilon_t \tag{E.2}$$

$$\begin{aligned}\Delta r_t = {} & \beta_0 + \sum_{j=1}^{q} \psi_j \Delta r_{t-j} + \sum_{k=0}^{p} \beta_{1+k} \Delta m_{1-k} + \sum_{i=1}^{n}\sum_{j=0}^{k} \gamma_{ij} \Delta\left(z_{t-j} m_{t-j}\right) + \sum_{i=1}^{n}\sum_{j=0}^{l} \pi_{ij} \Delta z_{t-j} \\ & + \eta_r r_{t-1} + \eta_m m_{t-1} + \sum_{i=1}^{n} \eta_{m,zi} \left(z_{i,t} m_{t-1}\right) + \sum_{i=1}^{n} \eta_{zi} z_{i,t-1} + \varepsilon_t\end{aligned} \tag{E.3}$$

$$\begin{aligned}\Delta r_t = {} & \beta_0 + \sum_{j=1}^{q} \psi_j \Delta r_{t-j} + \sum_{k=0}^{p} \beta_{1+K}^{+} \Delta m_{t-k}^{+} + \sum_{k=0}^{p} \beta_{1+K}^{-} \Delta m_{t-k}^{-} + \sum_{j=1}^{n}\sum_{i=0}^{k} \gamma_{ij} \Delta\left(z_{t-i} m_{t-i}\right) + \sum_{j=1}^{n}\sum_{i=0}^{l} \pi_{ij} \Delta z_t \\ & + \eta_r r_{t-1} + \eta_m^{+} m_{t-1}^{+} + \eta_m^{-} m_{t-1}^{-} + \sum_{i=1}^{n} \eta_{m,zi} \left(z_{i,t} m_{t-1}\right) + \sum_{i=1}^{n} \eta_{zi} z_{i,t-1} + \varepsilon_t\end{aligned} \tag{E.4}$$

The cointegration test consists of testing the following hypotheses, respectively:

- Equation (D.1): $H_0 : \eta_0 = \eta_1 = 0$ vs. $H_1 : \eta_0 \neq \eta_1 \neq 0$
- Equation (D.2): $H_0 : \eta_0 = \eta_1^{+} = \eta_1^{-} = 0$ vs. $H_1 : \eta_0 \neq \eta_1^{+} \neq \eta_1^{-} \neq 0$
- Equation (D.3): $H_0 : \eta_r = \eta_m = \eta_{m,z1} = \eta_{m,z2} = \ldots = \eta_{m,zn} = \eta_{z1} = \eta_{z2} \ldots \eta_{zn} = 0$ vs.

  $H_1 : \eta_r \neq \eta_m \neq \eta_{m,z1} \neq \eta_{m,z2} \neq \ldots \neq \eta_{m,zn} \neq \eta_{z1} \neq \eta_{z2} \ldots \eta_{zn} \neq 0$
- Equation (D.4): $H_0 : \eta_r = \eta_m^{+} = \eta_m^{-} = \eta_{m,z1} = \eta_{m,z2} = \ldots = \eta_{m,zn} = \eta_{z1} = \eta_{z2} \ldots \eta_{zn} = 0$ vs.

  $H_1 : \eta_r \neq \eta_m^{+} \neq \eta_m^{-} \neq \eta_{m,z1} \neq \eta_{m,z2} \neq \ldots \neq \eta_{m,zn} \neq \eta_{z1} \neq \eta_{z2} \ldots \eta_{zn} \neq 0$

In the bounds test, the null hypothesis is that no long-run relationship exists, i.e., the dependent and the independent variables do not move together across time in the long run. To this end, F-statistics are used to determine the global significance of the model, with particular attention on the lagged variables. Pesaran et al. (2001) supply bounds on the critical values for the F-test; however, because the I(0) and I(1) critical values generated by Pesaran et al. (2001)

are for large samples, I also use simulated values from Narayan (2005) who have generated critical values specifically for small sample sizes. The lower bound is based on the assumption that all variables are I(0), and the upper bound is based on the assumption that all variables are I(1). If the computed F-statistic falls below the lower bound, one can conclude that the variables are not cointegrated. If the F-statistic exceeds the upper bound, we find that we have cointegration. Finally, the test is inconclusive in the case where the F-statistic falls within the bounds. When the presence of cointegration is confirmed, the examination of both the short run and the long-run market-to-bank IRPT is performed using the error-correction framework.

## Appendix F: Test for Structural Breaks

Table F: Chow breakpoint test

| *Breakpoint* | *Statistic test* | | *p-value* |
|---|---|---|---|
| October 2005 | F-statistic | 18.12*** | 0.000 |
| | Log-likelihood ratio | 30.44*** | 0.000 |
| | Wald statistic | 36.23*** | 0.000 |
| August 2011 | F-statistic | 0.20 | 0.823 |
| | Log-likelihood ratio | 0.41 | 0.813 |
| | Wald statistic | 0.39 | 0.823 |
| June 2015 | F-statistic | 1.12 | 0.332 |
| | Log-likelihood ratio | 2.35 | 0.309 |
| | Wald statistic | 2.245 | 0.325 |

Notes: **** indicates rejection of the null hypothesis of no existence of a breakpoint at the 1% significance level*

## Appendix G: Results of the Cumulative Sum test (CUSUM test)

The CUSUM test (Page, 1954), which is based on the residuals from the recursive estimates, is a powerful tool for the detection of unknown structural breaks in the regression equation (Turner, 2010). To understand the CUSUM test procedure, consider $e_{t+1,t}$ as the one-step-ahead forecast error associated with forecasting $lr_t$ based on equations (E1), (E2), (E3), and (E4) for the period 2000:1–2016:3, thus $e_{t+1,t} = lr_t - lr_{t+1,t}$, which are the recursive residuals.

We also denote $\sigma_{1,t}$ as the standard error of the one-step-ahead forecast of $lr_t$ formed at time $t$, i.e. $\sigma_{1,t} = \sqrt{\operatorname{var}\left(e_{t+1,t}\right)}$. Now the standard recursive residuals ($w_{t+1,t}$) are defined as $w_{t+1,t} = e_{t+1} / \sigma_{1,t}$, where $w_{t+1,t} \sim i.i.d.\ N(0,1)$, and the CUSUM statistics are defined according to $CUSUM_t = \sum_{i=k}^{t} w_{i+1,i}$, $t = k, k+1, \ldots T-1$.

Under the null hypothesis, the $CUSUM_t$ statistic is drawn from a $CUSUM_{t-k}$ distribution, which is a symmetric distribution centred at 0. Its dispersion increases as the $t-k$ distribution increases. The CUSUM test is performed by constructing and analysing a CUSUM chart. The results of the CUSUM test for equations (E1)–(E4) are shown in Figures G.1–G.4. The null hypothesis is rejected at the 5% significance level.

Figure G.1: Simple linear model (Equation (1))

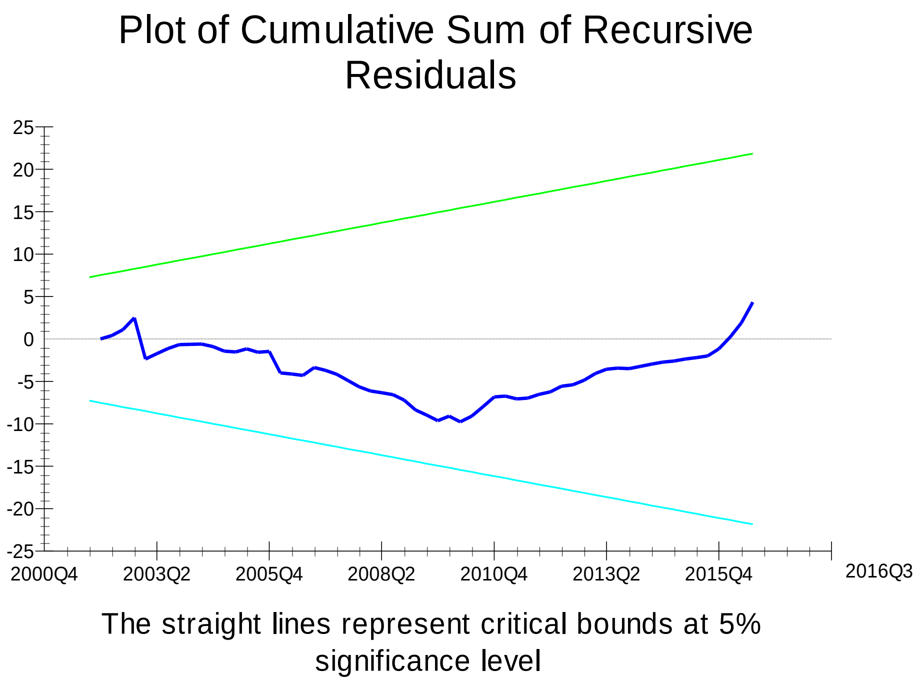



Figure G.2: Simple asymmetric model (Equation (3))

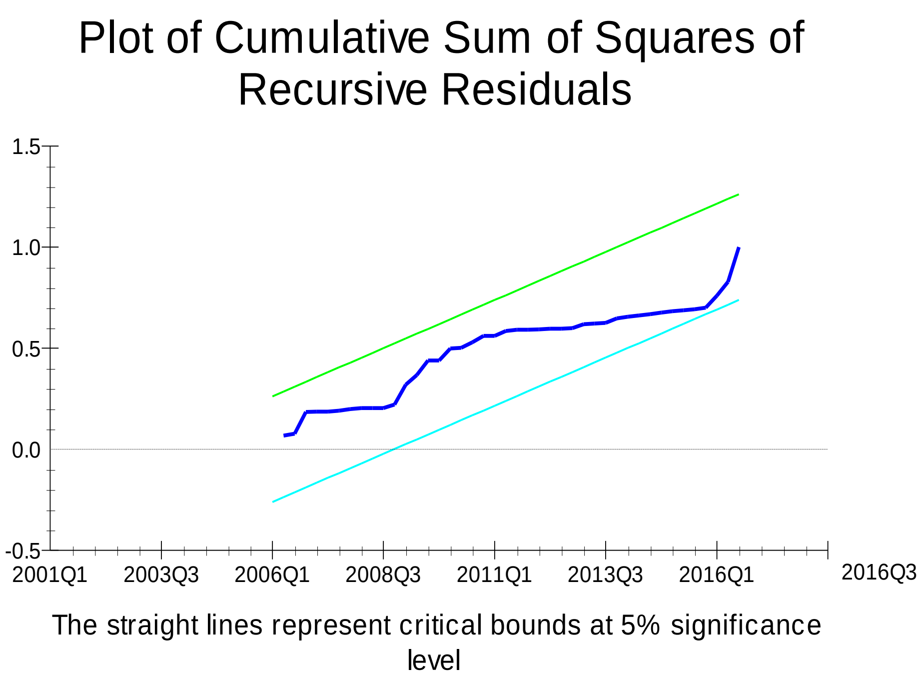
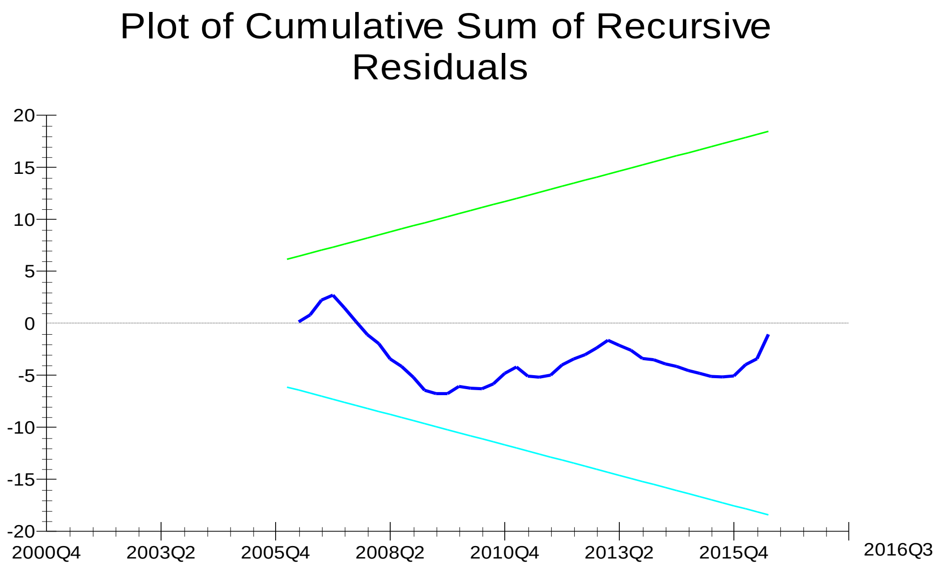
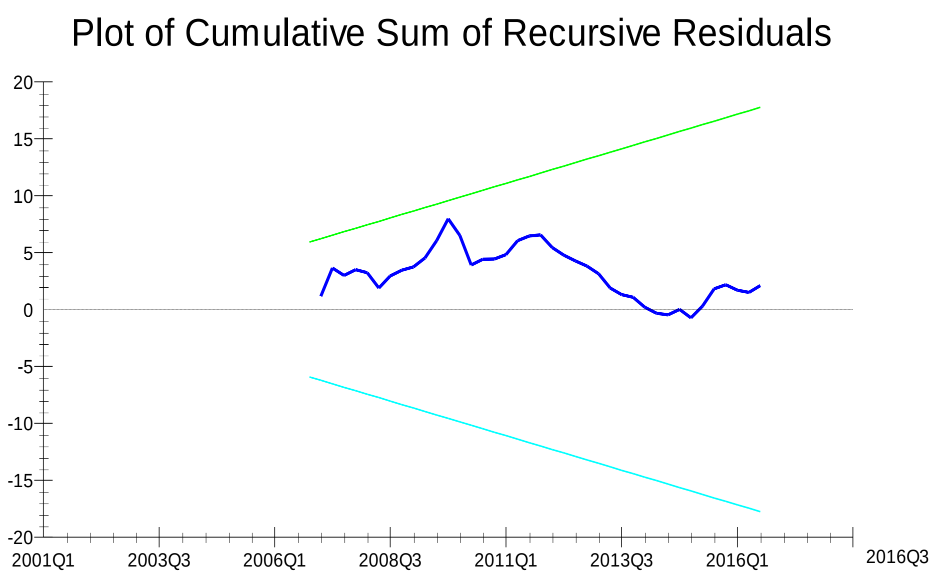


Figure G.3: Linear model with macroeconomic variables (Equation (6))




The straight lines represent critical bounds at 5% significance level

Figure G.4: Asymmetric model with macroeconomic variables (Equation (8))




The straight lines represent critical bounds at 5% significance level