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Interest Rate Pass-Through and Asymmetries in Retail Deposit and Lending Rates: An Analysis using Data from Colombian Banks

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Abstract

Using a sample of Colombian banks, we examine retail interest rate adjustment in response to changes in wholesale interest rates. Interest rate pass-through running from wholesale to retail rates is found to be both partial and heterogeneous across banks. This suggests that the effectiveness of monetary policy is limited. Further investigation reveals that the behaviour of retail deposit rates appears consistent with collusive behaviour between banks insofar as interest rates are more rapidly adjusted downwards than upwards. In the case of retail lending rates, it appears that banks more rapidly reduce than increase rates. This suggests that expansionary monetary policy in Colombia may be relatively more effective than contractionary policy.

Key words

interest rate pass-through

asymmetries

M/TAR model

JEL Classification

C33; E43

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1 Introduction

The behaviour of interest rates on deposits and loans in the retail banking sector has an important bearing on investment decisions and the real economy. While important drivers of movements in retail bank rates include the stance of monetary policy, the interbank or money market (wholesale) rate, there is considerable interest in how retail bank rates actually respond when such drivers change. In this respect, researchers have investigated the speed of response as well as the possibility that retail rates respond differently to increases and decreases in other interest rates. Following the seminal work by Stiglitz and Weiss (1981), interest rate rigidity can be attributed to adjustment costs and information asymmetries in credit markets. In a modelling approach based on an imperfectly competitive banking sector, Hannan and Berger (1991), Neumark and Sharpe (1992) and Freixas and Rochet (1997) suggest that both lending and deposit interest rates maintain a stable long-run equilibrium relationship with the interbank rate. Within this framework, error correction based on the adjustment of retail interest rates ensures that deviations from long-run equilibrium based on shocks are temporary. However, Hannan and Berger (1991) and Neumark and Sharpe (1992) advocate the collusive pricing hypothesis and the consumer behaviour hypothesis as alternative explanations for the extent of interest rate pass through and the adjustment of lending rates to changes in policy rates. As noted by Payne (2007), a downward rigidity of lending rates can be attributed to the reluctance of banks to decrease lending rates in fear of disrupting collusive arrangements and/or the hesitation by consumers to change lenders due to switching costs. On the other hand, the reaction from customers to lending rate increases and/or the adverse selection problem faced by lenders in an increasing interest rate environment may translate into upward rigidity in lending rates. This reasoning can be expanded to interest rates on bank deposits with banks more partial to raising rather than reducing deposit rates under the consumer behaviour hypothesis.

Numerous studies have examined the extent and nature of interest rate pass through and asymmetric adjustment based on various country samples. While much of this work is on the adjustment of retail interest rates with respect to changes in policy interest rates, there is a general finding that asymmetries are present with respect to how bank retail rates are adjusted. In terms of the relatively recent work, Payne (2007), using a momentum-threshold autoregressive (MTAR) specification, finds that the respective adjustable rate mortgages in the US and the federal funds rate are cointegrated, but with incomplete interest rate pass through. The results also indicate asymmetries in the response of the adjustable rates to changes in the federal funds rate. In an analysis of interest rates in Turkey, Yuksel and Ozcan (2013) employ the asymmetric threshold autoregressive (TAR) and MTAR procedures over the period December 2001 to April 2011 and find significant and complete pass-through between policy rate and loan rates. Positive and negative departures from the equilibrium converge to long-run path almost at the same speed. However, their analysis revealed that there is no significant relationship between policy rate and bank deposit rates due to sluggish adjustment of the latter. Zulkhibri (2012) examines Malaysia and finds that the pass through from money market rates to retail deposit and lending rates is incomplete. In addition to this, interest rate adjustments are found to be asymmetric, with more significant adjustments taking place under monetary easing than under monetary tightening. Wang and Thi (2010) find robust evidence that there exist the upward rigidity in the deposit rate and the downward rigidity in the lending rate in both Taiwan and Hong Kong. This is a finding that is consistent with the hypothesis of the collusive pricing arrangements.

In the case of Portugal, Rocha (2012) uncovers heterogeneous adjustments of bank rates as between sectors, between loans and deposits, and across maturities - which include complete long-run pass-through to corporate lending rates but rigidities for the personal sector. Verheyen (2013) presents results that point to considerable asymmetries especially with regard to the long-run pass-through of money market rate changes as well as some heterogeneity between EMU countries; see also Marotta (2009) and Bernhofer and van Treeck (2013). Outside of the Eurozone, Becker et al. (2012) find the presence of substantial asymmetries when exploring the pass-through of the official rate to the money market rate and of the market rate to the mortgage rate in the UK. Hofmann and Mizen (2004) look at deposit and mortgage products offered by individual UK financial institutions and find that the speed of adjustment in retail rates depends on whether the perceived 'gap' between retail and base rates is widening or narrowing. Further work on mortgage rates includes Valadkhani and Anwar (2012) who find that the Reserve Bank of Australia's rate rises have a much larger and more instantaneous impact on the mortgage rate than rate cuts. In contrast, Payne (2006) finds that US mortgage rates respond symmetrically to the federal funds rate in the long-run adjustment process. Lastly, Sznajderska (2013) examines the evidence for asymmetric effects in the Polish interest rate pass-through using TAR and MTAR models over the periods 2004m1- 2008m8 and 2004m1-2012m4. The results indicate that there are many more cases of asymmetric cointegration in the earlier shorter sample period than in the later longer sample period. Sznajderska suggests that the absence of cointegration may be due to disturbances in 2004, possibly connected with Poland's entry to the European Union, or also perhaps associated with methodological changes in the way interest rates are calculated and collected (in order to make them comparable with other EU countries).

In this paper, we estimate M/TAR models to examine if there is evidence of asymmetric behaviour in response to market conditions in the case of retail deposit and lending rates set by Colombian banks. While this is not the first case study of interest rate pass through in Colombia, we believe that the Colombian experience is indeed interesting on the grounds that it is an economy which about two decades ago embarked on a series of financial sector reforms aimed at increasing the degree of competition in retail banking. However, a cursory look of some recent basic banking sector indicators, such as the evolution of the number of banks and the associated five-bank concentration ratio (based on assets), actually reveals an increase in market concentration. This clearly makes it an open question as to which hypothesis (that is, collusive pricing versus consumer behaviour) is likely to prevail. In an earlier study, Iregui et al. (2002) offer an initial examination of the effects of the financial liberalisation measures on the dynamics of Colombian retail interest rates. These authors characterise the behaviour of aggregate deposit and lending rates in terms of the so-called smooth transition autoregressive models (STAR), and find evidence of relatively greater rigidity for both deposit and lending rate decreases when compared with increases.[1]

The distinguishing feature of our empirical analysis is that we take advantage of a highly disaggregated database that consists of the deposit and lending rates applied by the individual banks that comprise the Colombian banking sector. This is in marked contrast to much of the existing literature which, as indicated above, proceeds at a more aggregated sector-level basis.[2] The advantage of having data with such a level of disaggregation in our present study is that it is possible to determine whether the pass-through from market to retail interest rates as well as the presence of asymmetries (if any) is homogeneous across the banking institutions, or if they are dependent on the size of the intermediaries.

A further dimension that makes the Colombian experience attractive is

[1]It is worth mentioning that, similar to Iregui et al. (2002), we started our empirical analysis by fitting logistic and exponential STAR models to the data. However, the results (not reported here) indicated that the parameter which measures the smoothness in the value of the transition function was rather high, implying an almost instantaneous transition from one regime to the other. Bearing in mind that as the speed of transition tends to infinity the LSTAR and ESTAR models converge to the M/TAR model, the latter specifications were preferred over the former.

[2]An exception includes the recent paper by Valadkhani and Worthington (2014), who investigate asymmetric behaviour of Australia's Big-4 banks in the mortgage market.

that the effects of the Global Financial Crisis (GFC) have not been regarded as serious as it has been for other economies, mainly because of favourable commodity price movements related to Colombian exports and the country's strong economic policy framework.[3] This suggests that a zero lower bound on interest rates has been a less acute issue when it comes to lax monetary conditions. Instead, the existence of usury laws gives rise to an upper boundary on lending rates, which is commonly referred to as the interest rate of usury. This interest rate opens up the interesting possibility of an upper bound, which is in sharp contrast to most other countries which have had concerns with a lower zero bound during the GFC era.

The paper is organised as follows. Section 2 presents a brief overview of the developments in the Colombian financial sector, including an outline of the measures taken to increase competitiveness. Section 3 describes the econometric modelling strategy which is based on a threshold autoregressive approach for looking at asymmetric error correction towards long-run equilibrium between retail interest rates and interbank rates. Section 4 presents the data and the results of the empirical analysis. Our investigation reveals evidence that is consistent with the consumer reaction hypothesis being applicable in the case of lending rates, but the collusive behaviour hypothesis being applicable in the case of deposit rates. Section 5 offers concluding remarks.

2 An overview of recent developments in the Colombian banking sector

Until the late 1980s the banking sector in Colombia was subject to important restrictions in the form of high reserve requirements and liquidity ratios, controls on interest rates, and direct credit to specific sectors of the economy at subsidised interest rates. In this environment of "financial repression",

[3]See, e.g., the IMF country report 14/141.

government-owned banks held approximately 43% of the assets of the financial sector (about 20% of GDP), while banks with foreign participation amounted to only 3% of the total assets of the financial system (approximately 1% of GDP); see Uribe and Vargas (2003). In addition, Clavijo et al. (2006) observe that the asset side of the bank's balance sheet was rather specialised, as these institutions mostly dealt with specific sectors of the economy, such as agriculture, construction, industry, and commerce, among others.

In the early 1990s, the Colombian government embarked on a major programme of reforms aimed at liberalising the economy with the purpose of making it more competitive. The reforms comprised several fronts: trade relations, foreign exchange regime, labour market, social security, government finances and the financial sector. Focussing on the latter, Uribe and Vargas (2003) and Arango (2006) summarise the package of financial liberalisation reforms in four main laws. Among other measures, Law 45 of 1990 redefined the structure of the financial sector; relaxed the requirements for entry and exit of intermediaries; regulated mergers, acquisitions and liquidations; substantially reduced reserve requirements; and liberalised interest rates, although ceilings regulated with the existence of usury laws are still in place today in some segments of the market.[4] Law 9 of 1991 replaced the foreign exchange regime that existed at the time for another one based upon the principle of equal treatment between domestic and foreign investors, so that they would have access to all sectors of the economy including, of course, the financial system. Law 31 of 1991, following precepts established in the Constitutional Reform of 1991, strengthened the institutional structure of the

[4]Currently, the Colombian Superintendency of Financial Institutions has two categories of interest rate caps, namely the usury interest rate for conventional (i.e. ordinary and consumption) loans and the usury interest rate for microcredit (i.e. loans to microenterprises). Supervised institutions must report their rates by category on a weekly basis, and these data are used by the Superintendency to fix the level of the usury rate for conventional loans and microcredit quarterly and annually, respectively. In both cases, the usury rate is set as 1.5 times the corresponding average rate for the system as a whole.

economy by creating a central bank independent from the government's executive branch. Lastly, Law 35 of 1993 enhanced the spectrum of operations that the different intermediaries could undertake, facilitating the transition towards a system of "universal" banking.

As a result of the package of reforms summarised above, by December 2010 the share of government-owned banks in the assets of the financial sector had decreased to approximately 5% (about 3% of GDP), while that of banks with foreign participation had increased to approximately 18% (about 9% of GDP). As to the evolution through time of the number of banks, it went up from 36 in 1985 to 41 in 1997; see Appendix 1 in Arango (2006). In the late 1990s, a period during which the Colombian economy witnessed the deepest recession ever recorded, the number of commercial banks fell as a result of liquidations, acquisitions and mergers. By 2002, the number of banks was reduced to 29, and by 2010 this figure had reduced even further to 18 institutions. In terms of the state of competitive conditions prevailing in the banking sector, it appears that market concentration has increased in recent years. Indeed, while in 2001 five banks accounted for as much as 43% of the total assets of the banking sector, in 2010 this percentage had risen to approximately 63%. At this point it is important to observe that the policy developments summarised above can be put in the context of the empirical analysis of the following sections. Indeed, if the above mentioned reforms succeeded in making the banking sector more competitive, then we would expect stronger support for the consumer reaction hypothesis rather than the collusion pricing hypothesis. Associated with this, there is also the question of whether support for either hypothesis is generalised throughout the banking sector or focalised in some specific institutions within it.

The above discussion pays special attention to recent initiatives that took place in Colombia. In disentangling the potentially different effects of institutional changes from environmental changes, it should be noted that Colombia was relatively less affected by the GFC compared to other countries. Further to this, the Colombian growth performance over our 2002-14 study period has been impressive when benchmarked against major economies making Colombia the third largest Latin American economy. In addition, the Colombian economy has seen a reduced public deficit combined with strong confidence. Despite the continued reliance on commodity exports, such factors have served to provide a relatively stable and conducive environment in which Colombian banks have been able to operate effectively and develop as a result of the institutional reforms. While such environmental changes have influenced the amount of business conducted by Colombian banks, the institutional changes have influenced the way in which banks conduct their activities.

3 Econometric modelling strategy

We are interested in testing whether there is a long-run equilibrium relationship between the non-stationary retail interest rate of a bank, say y_t, and the non-stationary market interest rate, say x_t, focusing on whether the residuals u_t that result from estimating the bivariate relationship:

$$y_t = \beta_1 + \beta_2 x_t + u_t, \tag{1}$$

are stationary, also denoted $\sim I(0)$ for short. In the well-known Engle and Granger (1987) approach, testing for cointegration involves using the OLS residuals from (1) to estimate:

$$\Delta u_t = \rho u_{t-1} + \varepsilon_t, \tag{2}$$

where up to p lags of the dependent variable can be included in the right-hand side to guarantee that the error term ε_t is a white-noise disturbance. Under the null hypothesis of no cointegration $\rho = 0$, while under the alternative of cointegration $\rho < 0$. Because of the linearity in (2), evidence in favour of the alternative hypothesis implies symmetric adjustment insofar negative and positive residuals give rise to the same speed of adjustment back to long-run equilibrium. Within this framework, Enders and Siklos (2001) indicate that the model in (2) would be incorrectly specified if adjustment rather occurs in an asymmetric fashion, as implied for instance in the so-called threshold autoregressive (TAR) model, written as:

$$\Delta u_t = \rho^+ I_t u_{t-1} + \rho^- I_t u_{t-1} + \varepsilon_t, \tag{3}$$

where I_t is an indicator function that is equal to one if $u_{t-1} \geq \tau$, zero otherwise, and τ is the value of the threshold parameter. In terms of asymmetric adjustment, finding that ρ^+ is further from zero than ρ^- would imply that retail interest rates are likely to respond relatively more quickly to decreases in the level of market interest rates than increases. That is, there is a different speed of adjustment back towards long-run equilibrium based on loose or tight monetary policy. In addition to this, Enders and Siklos (2001), following Enders and Granger (1998) among others, also suggest an alternative adjustment specification in which the indicator function I_t does not depend on the level of residual series but on its first difference, so that I_t is equal to one if $\Delta u_{t-1} \geq \tau$, and zero otherwise. This alternative adjustment process results in the so-called momentum threshold autoregressive model, or MTAR model for short. The interpretation of asymmetry in the MTAR model alters insofar as when ρ^+ is further from zero than ρ^- would now imply that retail interest rates are likely to respond relatively more quickly to decreases in the change of market interest rates than increases in the change. This time there is a different speed of adjustment based on looser or tighter monetary policy. In both types of model, an examination of ρ^+ and ρ^- enables us to assess the relevance of the collusive pricing and consumer behaviour hypotheses in explaining bank behaviour. In the case of deposit rates, ρ^+ more negative than ρ^- might indicate behaviour that is consistent with the collusive pricing hypothesis because banks appear to adjust their deposits relatively more quickly in response to loose or looser monetary policy compared to the tight or tighter monetary policy. On the other hand, in the case of lending rates ρ^+ more negative than ρ^- might indicate behaviour consistent with the consumer behaviour hypotheses because banks appear to reduce their lending rates relatively more quickly when monetary policy is loose or looser.

Enders and Siklos (2001) observe that regardless of whether one is assuming a TAR or a MTAR type of adjustment, the Engle-Granger setup is a special case of the model in (3) when $\rho^+ = \rho^-$. In addition, they indicate that although in many economic applications setting $\tau = 0$ seems a natural way to proceed, the value of this parameter is in general unknown, so that it has to be estimated along with the values of ρ^+ and ρ^-. Enders and Siklos (2001) recommend the use of the consistent threshold estimation method advocated by Chan (1993) which involves the following steps. First, the estimated residual series u_t is sorted in ascending order. Second, the highest and lowest 15% of the sorted series are excluded so as to ensure an adequate number of observations on each side of the threshold. As our database consists of 142 time observations, this leaves 99 potential threshold values to be considered. Third, these threshold values are used to create 99 indicator functions I_t, which are subsequently used to estimate the same number of TAR models as given in (3). For each of these equations one records the residual sum of squares, and the equation yielding the smallest of these is deemed to be the appropriate consistent estimate of the threshold parameter.

Testing for cointegration in the presence of threshold adjustment involves testing the joint null hypothesis that $\rho^+ = \rho^- = 0$, and the resulting testing statistic is denoted Φ when $\tau = 0$ or Φ^* when the threshold is estimated as outlined in the previous paragraph. The corresponding critical values are reported in Tables 1 and 5 of Enders and Siklos (2001), respectively. These authors also proposed a t-Max statistic, calculated as the largest of the two individual t statistics to test the hypotheses that $\rho^+ = 0$ and $\rho^- = 0$, although this is rarely applied in practice because it exhibits low power. If one does not have a specific preference for either the TAR or the MTAR specification, then it is often recommended to estimate both models and choose the one with, for instance, the lowest residual sum of squares as the preferred specification. In addition to cointegration with threshold adjustment, Enders and Siklos (2001) also suggest testing the null of symmetric adjustment which is postulated as $\rho^+ = \rho^-$ and can be tested using a standard F distribution.

4 Data and empirical analysis

This paper examines wholesale and retail interest rates in Colombia. The data used in the empirical analysis correspond to monthly interest rate data (expressed in percentage terms) from October 2002 to July 2014, for a total of 142 time observations. The wholesale (or money market) interest rate is the interbank rate series, denoted ib_t, is taken from the Colombian central bank (Banco de la República).[5] Retail interest rates are individual deposit and lending rates offered by 15 banking institutions. In December 2013, these banks collectively account for almost the totality of the assets of the banking sector (that is, 97.3%), and more than half of the assets of the supervised financial system (that is, 54.5%). The representativeness of the selected sample of banks on the aggregate Colombian economy is well exemplified by the fact that their assets represent about 55% of GDP in 2013. The choice of banks is dictated by the desire to assemble a consistent database over the largest possible study period given data availability. The deposit (lending) rates are weighted averages computed over all maturities. Thus, the deposit

[5] According to the central bank (see *www.banrep.gov.co/es/tib*), the interbank rate refers to the price of the operations in domestic currency that are undertaken between financial intermediaries to solve liquidity problems overnight (although the term of the rate is typically one day, it can vary during weekends or when there are holidays during weekdays). Loans between intermediaries do not require collateral, and so the interbank rate reflects the credit risk associated to the parts involved in the operations. Moreover, the level of the rate reflects liquidity conditions in the money market, and so it is a good indicator of the stance of monetary policy in Colombia.

rate of bank i, denoted d_{it}, is the weighted average rate paid by the bank on 30-, 45-, 60-, 90-, 120-, 180-, 360-, and more than 360-day time deposits (CDTs), using the amount of deposits at each maturity as weight. In turn, the lending rate of bank i, denoted l_{it}, is the weighted average rate charged on 1-, 3- and 5-year consumption loans, using the amount of loans at each maturity as weight.

Retail interest rates are obtained from weekly surveys on deposit and lending rates reported by the banks to the Colombian Superintendency of Financial Institutions. These surveys also report data on the amount of deposits and loans for the different maturities. The index $i = 1, ..., 15$ denotes the banks under consideration. To ease the interpretation of the results, hereafter banks are reported in descending order according to the size of their assets; thus, d_{1t} (l_{1t}) denotes the deposit (lending) rate of the bank with the largest assets while d_{15t} (l_{15t}) denotes the corresponding rate of the bank with the smallest assets in the sample.[6] Based on the data, Figure 1 shows that during the sample period under consideration the average range of variation of the deposit interest rates, calculated as the difference between the maximum and minimum interest rates, is 2.5 percentage points as opposed to 13.2 percentage points for the lending interest rate. The figure also shows the evolution of the interest rate of usury, which constitutes the maximum interest rate that the supervised financial sector can charge without incurring in the (illegal) practice of usury. As can be seen from the figure, during the first 10 years or so of the sample period the maximum lending interest rate was very close to this ceiling, while distancing from it during the last years of the period under consideration. Interestingly, the detachment between the lending and the usury interest rates coincides with a narrowing of the range of variation of deposit interest rates. This graphical evidence appears

[6] An alternative criterion to sort banks could well be based on their riskiness as measured, for example, by their credit ratings. However, we do not have data on bank credit ratings and so this alternative sorting is not carried out.

to suggest that the behaviour of banks differs depending on whether they are acting as borrowers or lenders of funds. We shall return to this feature of the interest rate data when discussing the results of our econometric analysis.

The empirical analysis starts off by examining the time series properties of the interest rates under consideration. For this, we employ both the ADF and ADFmax unit root tests of Dickey and Fuller (1979) and Leybourne (1995), respectively. The former is the well-known regression-based procedure that, in its basic form, tests the null that the first order autoregressive coefficient in a first order autoregressive equation is equal to one (where this simple specification can be suitably generalised to allow for the possibility of deterministic components in the series and/or with the presence of residual serial correlation). The latter is computed as the maximum statistic that results from applying the ADF test to both the forward and reversed realisations of the data. Table 1 summarises the results of the unit root tests, when the test regression includes an intercept and four lags to account for residual serial correlation. The results in this table support the view that all the interest rates can be best described as non-stationary processes, as the null hypothesis is not rejected at the 5% significance level. Qualitatively similar findings are obtained when the test regression also includes a trend and/or when the number of lags is determined using a data-dependent procedure such as the Akaike information criterion.

Once the order of integration of the interest rates has been determined, we proceed with the cointegration analysis for which we commence by estimating the long-run Engle-Granger cointegration equation using OLS, the results of which are reported in Table 2. The importance of the cointegrating regression relies on the fact that the slope coefficient is commonly interpreted as the long-run pass-through coefficient from the market rate to the deposit (lending) rate. As can be seen in Table 2, the estimate of this coefficient reveals a large degree of heterogeneity across all banks. Indeed, our findings indicate that for deposit rates the long-run pass-through coefficient ranges from 0.44 to 0.86 (on average 0.70), compared to a much wider range of variation from 0.06 to 1.26 for lending rates (on average 0.81). Given that most of these slopes are positive but below unity, these results might initially suggest the presence of weak pass through as an increase of the market rate of one percentage point leads to a less than proportional increase in the deposit (lending) rate.

These average pass through coefficients are higher than those reported in the studies such as Payne (2007) who finds that the pass through coefficient between Federal Funds and Mortgage rates is of the order 0.5, and Zulkhibri (2012) who computes a mean long-run pass through coefficients of 0.66 across 14 Malaysian commercial banks and finance companies. In contrast, our average estimates for Colombia are below those provided by Valadkhani (2013) who reports pass-through coefficients greater than unity for some Australian financial institutions, but our findings are instead comparable to Valadkhani and Worthington (2014) who also find that the pass through coefficients that link mortgage rates to the official cash rate for the Big-4 Australian banks are of the order 0.7-0.8. This in turn reflects on the long-run effectiveness of monetary policy. Incomplete pass-through means that effectiveness is somewhat blunted, but it is notable that pass through to lending rates is greater than for deposit rates. In further investigation of Colombian data, we also explored whether an examination of other rates might provide additional insights into retail bank behaviour. For example, the employment of credit card interest rate data in place of lending rates still provided a cointegrating equation indicative of incomplete pass-through with a smaller average long-run pass-through coefficient of 0.45. This lower degree of pass-through is a reflection of credit card interest rates that have typically followed and remained slightly below the usury rate over the whole sample period.

However, the validity of the results reported in Table 2 is of course subject to the existence of a long-run equilibrium relationship between the market rate and the deposit (lending) rate as otherwise the regression results would be spurious. Table 3 reports the results of the cointegration analysis. The testing strategy that we adopt follows a general-to-specific approach in the sense that we start off by testing for threshold cointegration for both the TAR and MTAR specifications, where the model with the smallest residual sum of squares is selected as the preferred specification. Next, we consider the possibility of linear cointegration if the null hypothesis of symmetry in the selected model, that is $H_o : \rho^+ = \rho^-$, is not rejected.

The top half of Table 3 reveals that the TAR model is chosen for the deposit rates of six banks while the MTAR specification is the preferred option for six of the remaining nine banks. Focusing first on the six TAR models, in all cases we find support for the presence of cointegration with threshold adjustment as the Φ^* statistics reject the null hypothesis at least at the 90% significance level. Furthermore, the F statistic to test the null hypothesis of symmetry is clearly rejected in all six models with ρ^+ being greater (in absolute value) than ρ^-, suggesting that the speed of adjustment back to equilibrium is quicker when deposit rates are above the value implied by the long-run relationship with ib_t. Regarding the nine other banks, there are two deposit rate series (for banks 4 and 6) for which we fail to reject the null hypothesis that there is no cointegration with threshold adjustment, and three deposit rates (for banks 5, 10 and 15) for which there is evidence of cointegration but where adjustment appears to be symmetric, as the hypothesis that $\rho^+ = \rho^-$ is not rejected (this conclusion is subsequently verified by performing the Engle-Granger cointegration analysis). For banks 3, 7, 11 and 14 there is support for cointegration with momentum threshold adjustment. Also, with the exception of bank 3, adjustment to the long-run level occurs more quickly when deposit rates are above equilibrium. In short, for deposit rates there is evidence of cointegration, with or without (momentum) threshold adjustment, in 10 out of 15 cases.

Turning to the results for lending rates, we uncover evidence of cointegration with threshold adjustment for seven banks. Of these, the MTAR type of adjustment appears to be the preferred specification in six out of the seven cases. In general terms, adjustment appears to occur more quickly when lending rates are above their long-run equilibrium, although such behaviour is not supported by the data in the cases of banks 5 and 10.[7] In summary, the results of the cointegration analysis suggest that the idea of the existence of a long-run equilibrium relationship between the market interest rate and retail rates appears to be more prevalent for deposit rates than for lending rates. Given that the long-run effectiveness of monetary policy hinges on the extent to which retail borrowing costs respond to wholesale rates, these results are providing a picture of limited policy effectiveness.

Having a majority of cases where the significant error correction coefficient is more negative above the threshold than below when it comes to either deposit or lending rates implies that Colombian banks (at least during the sample period under analysis) are quicker to reduce their retail interest rates than increase them when it comes to equal and be judged opposite changes in ib_t. Interestingly, these findings differ from those obtained by Iregui et al. (2002) who, using aggregate data for the 1990s and early 2000s, report evidence of greater rigidity for both deposit and lending rate decreases. In terms of the two hypotheses of interest, our findings are thus somewhat mixed. Indeed, while the consumer reaction hypothesis rather than the collusion hypothesis might be generally the more applicable in the case of lending rates, the collusion pricing hypothesis appears to be generally the more applicable one in the case of deposit rates. With the absence of a binding zero lower bound resulting from a limited impact from the GFC on Colombia, retail banks have been quicker and able to reduce their deposit rates in response to monetary loosening. These findings perhaps point towards the relative

[7] In the case of bank 10, however, these findings ought to be interpreted with caution because the estimated coefficient for negative deviations from equilibrium is positive and statistically different from zero, so that one of the stationarity conditions of the threshold autoregressive model is not satisfied.

success of the reforms in promoting increased competition in lending rather than deposit taking. Where asymmetries are confirmed, the MTAR model is preferred mostly in the case of the lending rate series. Given the nature of the MTAR threshold, the support for the consumer reaction hypothesis is based on evidence that lending rates exhibit a faster speed of adjustment back towards long-run equilibrium in an environment of falling rather than rising interbank rates.

With reference to the earlier abovementioned studies, Payne (2007) finds that adjustment towards the long-run equilibrium is faster when the US federal funds rate is falling relative to the adjustable rate. Zulkhibri (2012) reports that the adjustment in lending rates tends to be more sluggish than that of deposit rates. Valadkhani (2013) finds that the three largest domestic banks pass on the Reserve Bank of Australia's official rate rises to borrowers more than they do for rate cuts, affecting the efficacy of expansionary versus contractionary monetary policy. Further to this, Valadkhani and Worthington (2014) report that when mortgage rates are substantially above the equilibrium path, there is no significant attempt to lower rates, but faster adjustment occurs when rates are below equilibrium values. Sznajderska (2013) finds that Polish short-term deposit rates adjust significantly stronger to decreases than increases in the money market rate. In the case of Colombia, evidence that some banks more readily reduce lending rates than increase them in response to movements in wholesale rates, suggests that Colombian expansionary monetary policy is likely to be more effective than contractionary monetary policy. This perhaps is the opposite to what is implied by other country studies. However, it should be remembered that the majority of the sample does not exhibit asymmetries in the case of lending rates.

In the light of our findings, Figure 1 demonstrates the heterogeneity in bank behaviour that we have detected through the widening and narrowing of spreads in deposit and lending rates over time. One can indeed argue that the competition in lending rates has been facilitated by the rather limited range of variability in deposit rates across institutions. Our findings have pointed towards Colombian banks being quicker to reduce retail rates than increase them. From the mid-2000s, it can be seen that the spread in deposit rates appears to narrow when rates in general are falling. This is consistent with a greater general willingness on the part of banks to reduce rather than increase deposit rates. If we consider lending rates, then this pattern is also discernible to some extent in the mid- and late-2000s. Not only is the range of interest rate variation much smaller in the case of deposit rates, but Figure 1 also shows that while early in the sample period the maximum lending rate was pretty close to the interest rate of usury, during the last five years or so the former has consistently detached from the latter. Based on the presence of usury laws, a ceiling or upper bound on interest rates might lead one to expect faster adjustment from above when it comes to lending rates. However, the observation from Figure 1 of consistent detachment of lending rates from usury rates since the early part of the study period, may lend further support to the idea that competition has increased. This may account for the relative insignificance of ρ^+ that we have detected. From the practical point of view, our findings support the view that although depositors rarely benefit from attractive interest rates (as they all seem to be practically the same), there is scope for lenders to benefit from low rates if they look for bargains.

Lastly, we are now in position to gain some insight into both the short- and long-run dynamic behaviour of retail interest rates by formulating and estimating the underlying error correction models. The error correction models for the variations in the deposit and lending rates are summarised in Tables 4 and 5, respectively, where the significance of the error correction terms can be interpreted as an indication of long-run causality while that of the coefficients on ib_t and its lags is related to short-run causality. The results indicate that the effect of a one percentage point increase of ib_t on d_{it} ranges between 0.223 and 0.640 percentage points (in banks 7 and 11, respectively), and that it is statistically different from zero in all banks. Additionally, for some banks past changes in ib_t also have statistically significant effects. However, the overall picture is that short-run pass through is confirmed in the case of deposit rates albeit in an incomplete sense with short-run coefficients on ib_t that sum to less than unity. In sharp contrast to this behaviour, the only association between ib_t and l_{it} appears to exist for bank 15 which is the smallest bank in the sample, and with a coefficient whose magnitude (i.e. 1.515) suggests some sort of over adjustment. For the remaining six banks the estimated impact effect of ib_t of l_{it} is not statistically different from zero. Thus, ib_t affects l_{it} through its past changes and also through its presence in the positive and negative error correction terms lagged one period, denoted ect^+_{t-1} and ect^-_{t-1}, respectively.

5 Concluding remarks

It is important to further our understanding of how banks individually adjust their retail interest rates to changes in market conditions. In this paper, we have focussed on how Colombian banks have adjusted their deposit and lending interest rates in response to changes in the interbank (or market) rate. Colombia constitutes an interesting case study given the introduction of reforms aimed at achieving a more competitive banking sector. In addition to this, it is a country where, despite the Global Financial Crisis, the existence of usury laws actually gives rise to a greater likelihood of a binding upper bound on interest rates rather than a lower bound. In common with other studies based on aggregate data, our findings at the bank-level suggest that interest rate pass through is partial. However, in sharp contrast to the literature, we further find that asymmetries are present with respect to the adjustment of retail deposit interest rates in response to interbank rates. More specifically, we find deposit interest rate behaviour is consistent with collusive behaviour on the part of banks' retail activity insofar as deposit rates more readily adjusted downwards than upwards in response to changes in interbank rates. In the case of lending rates, there is less evidence of asymmetries across the banks, but when this evidence is found it is consistent with banks more readily reducing lending rates than increasing them. It would appear, therefore, that the success of the competitive reforms have had limited success in achieving their objectives. There is already evidence that market concentration has increased in the banking sector, and it is possible that asymmetric adjustment of lending rates is attributable to the presence of an interest rate ceiling based on usury laws.

Given that the long-run pass-through from wholesale interest rate into retail rates is incomplete, this serves to suggest that Colombian monetary policy effectiveness appears limited over the sample period under consideration. The evidence that some banks more readily reduce lending rates than increase them in response to movements in wholesale rates, suggests that Colombian expansionary monetary policy is likely to be more effective than contractionary monetary policy. For a country that was relatively unaffected by the global financial crisis, the presence of usury laws and an interest rate ceiling assumes a greater relevance than the perceived need for quantitative easing with potential constraints of a zero lower bound. However, it should be remembered that for many banks such asymmetries with respect to lending rates are not present. This suggests that an assessment of asymmetries in monetary policy effectiveness might play closer attention to the lending activities of particular banks in terms of which sectors of the Colombian economy they are most actively involved with.

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Table 1: Unit root tests

Series	ADF	p-value	ADFmax	p-value
ib	-2.162	[0.221]	-2.162	[0.088]
d_1	-2.028	[0.275]	-2.028	[0.116]
d_2	-1.950	[0.309]	-1.950	[0.135]
d_3	-1.596	[0.482]	-1.396	[0.336]
d_4	-1.346	[0.607]	-1.346	[0.359]
d_5	-1.819	[0.370]	-1.819	[0.173]
d_6	-1.697	[0.430]	-0.619	[0.707]
d_7	-1.156	[0.692]	-1.070	[0.493]
d_8	-1.818	[0.371]	-1.818	[0.173]
d_9	-1.347	[0.606]	-1.347	[0.358]
d_{10}	-1.648	[0.456]	-1.648	[0.231]
d_{11}	-1.775	[0.391]	-1.775	[0.187]
d_{12}	-1.618	[0.471]	-1.617	[0.243]
d_{13}	-1.649	[0.455]	-1.649	[0.231]
d_{14}	-1.633	[0.463]	-1.633	[0.237]
d_{15}	-1.886	[0.338]	-1.716	[0.206]
l_1	-2.125	[0.235]	-2.121	[0.096]
l_2	-1.616	[0.472]	-0.758	[0.645]
l_3	-2.396	[0.145]	-1.686	[0.217]
l_4	-1.615	[0.472]	-1.177	[0.440]
l_5	-0.776	[0.822]	-0.776	[0.636]
l_6	-1.938	[0.314]	-1.674	[0.221]
l_7	-0.837	[0.805]	-0.343	[0.810]
l_8	-2.504	[0.117]	-2.048	[0.112]
l_9	-1.369	[0.596]	-1.369	[0.348]
l_{10}	-0.317	[0.918]	-0.316	[0.819]
l_{11}	-2.331	[0.164]	-1.480	[0.299]
l_{12}	-2.053	[0.264]	-1.211	[0.423]
l_{13}	-1.664	[0.447]	-1.664	[0.225]
l_{14}	-2.387	[0.147]	-2.211	[0.079]
l_{15}	-2.174	[0.217]	-1.776	[0.186]

Note: The p-values for the ADF and ADFmax tests are based on response surfaces estimated by MacKinnon (1996) and Otero and Smith (2012), respectively.

Table 2: Cointegration equations

Series	Const.	(s.e.)	ib	(s.e.)	$\bar{R}^2$
d_1	0.528	(0.179)	0.742	(0.029)	0.826
d_2	1.442	(0.153)	0.688	(0.025)	0.848
d_3	0.917	(0.207)	0.696	(0.033)	0.757
d_4	0.837	(0.174)	0.672	(0.028)	0.804
d_5	0.412	(0.130)	0.806	(0.021)	0.913
d_6	1.590	(0.347)	0.438	(0.056)	0.302
d_7	1.564	(0.234)	0.717	(0.037)	0.722
d_8	0.668	(0.147)	0.783	(0.024)	0.886
d_9	1.710	(0.135)	0.747	(0.022)	0.894
d_{10}	0.942	(0.154)	0.864	(0.025)	0.897
d_{11}	0.663	(0.157)	0.775	(0.025)	0.871
d_{12}	1.263	(0.161)	0.564	(0.026)	0.772
d_{13}	1.919	(0.140)	0.608	(0.022)	0.838
d_{14}	1.231	(0.140)	0.731	(0.022)	0.883
d_{15}	1.913	(0.225)	0.711	(0.036)	0.733
l_1	11.043	(0.390)	0.871	(0.062)	0.579
l_2	10.527	(0.633)	0.914	(0.101)	0.363
l_3	18.642	(0.722)	0.457	(0.116)	0.094
l_4	14.087	(0.706)	1.020	(0.113)	0.363
l_5	9.936	(0.327)	1.132	(0.052)	0.767
l_6	11.942	(0.845)	0.694	(0.135)	0.152
l_7	14.460	(0.939)	1.261	(0.150)	0.329
l_8	22.079	(0.554)	0.055	(0.089)	-0.004
l_9	10.696	(0.508)	1.097	(0.081)	0.562
l_{10}	9.092	(0.341)	0.699	(0.055)	0.536
l_{11}	18.851	(0.637)	0.118	(0.102)	0.002
l_{12}	16.776	(0.605)	1.037	(0.097)	0.446
l_{13}	16.251	(0.416)	0.847	(0.067)	0.533
l_{14}	20.763	(0.573)	0.818	(0.092)	0.357
l_{15}	15.234	(0.717)	1.128	(0.115)	0.404

Table 3: Cointegration analysis

Series	Model	Φ^*	$\rho^+ = \rho^-$	p-val	t-stat	ρ^+	(s.e.)	ρ^-	(s.e.)
d_1	TAR	9.19^b	10.30	[0.002]		-0.166	(0.039)	-0.008	(0.031)
d_2	TAR	7.28^b	4.25	[0.041]		-0.160	(0.044)	-0.043	(0.036)
d_3	MTAR	9.75^c	6.24	[0.014]		-0.036	(0.025)	-0.142	(0.034)
d_4	MTAR	5.39	3.21	[0.076]		0.073	(0.101)	-0.118	(0.037)
d_5	Linear				$-4.61^\ddagger$	-0.184	(0.040)	-0.184	(0.040)
d_6	MTAR	4.88	4.81	[0.030]		-0.092	(0.031)	-0.015	(0.017)
d_7	MTAR	6.33^a	3.38	[0.068]		-0.197	(0.071)	-0.058	(0.026)
d_8	TAR	10.78^c	9.39	[0.003]		-0.214	(0.046)	-0.024	(0.042)
d_9	TAR	8.69^b	6.19	[0.014]		-0.191	(0.046)	-0.028	(0.047)
d_{10}	Linear				$-3.51^\dagger$	-0.093	(0.026)	-0.093	(0.026)
d_{11}	MTAR	10.91^c	11.00	[0.001]		-0.275	(0.059)	0.029	(0.073)
d_{12}	TAR	12.77^c	11.75	[0.001]		-0.145	(0.029)	-0.014	(0.025)
d_{13}	TAR	6.37^a	5.32	[0.023]		-0.164	(0.047)	-0.024	(0.040)
d_{14}	MTAR	11.01^c	4.88	[0.029]		-0.280	(0.074)	-0.100	(0.036)
d_{15}	Linear				$-3.74^\dagger$	-0.089	(0.024)	-0.089	(0.024)
l_1	TAR	4.58	5.31	[0.023]		-0.247	(0.083)	-0.032	(0.046)
l_2	TAR	8.42^b	7.97	[0.006]		-0.149	(0.036)	-0.003	(0.037)
l_3	MTAR	3.63	2.12	[0.148]		-0.101	(0.038)	-0.021	(0.040)
l_4	MTAR	5.60	4.77	[0.031]		-0.231	(0.075)	-0.048	(0.038)
l_5	MTAR	11.90^c	8.32	[0.005]		-0.200	(0.075)	-0.664	(0.151)
l_6	TAR	5.77	6.43	[0.012]		-0.029	(0.048)	-0.245	(0.073)
l_7	TAR	4.30	5.29	[0.023]		-0.008	(0.021)	-0.108	(0.037)
l_8	MTAR	6.07^a	6.06	[0.015]		-0.311	(0.092)	-0.057	(0.052)
l_9	MTAR	4.80	4.51	[0.036]		0.064	(0.095)	-0.160	(0.054)
l_{10}	MTAR	13.30^c	24.69	[0.000]		-0.238	(0.060)	0.267	(0.084)
l_{11}	MTAR	6.08^a	7.65	[0.007]		-0.147	(0.043)	0.024	(0.045)
l_{12}	MTAR	5.26	2.04	[0.156]		-0.271	(0.126)	-0.085	(0.035)
l_{13}	MTAR	10.55^c	8.83	[0.004]		-0.488	(0.107)	-0.101	(0.093)
l_{14}	TAR	5.62	7.24	[0.008]		-0.254	(0.076)	-0.024	(0.041)
l_{15}	MTAR	10.65^c	9.50	[0.003]		-0.364	(0.080)	-0.070	(0.058)

Note: The critical values for Φ^* are taken from Enders and Siklos (2001), Table 5 and $T = 100$. ^{a}, ^{b} and ^{c} indicate that the null hypothesis of no threshold cointegration is rejected at the 90, 95 and 99%, respectively. The column labelled $\rho^+ = \rho^-$ presents the symmetry test followed by the associated probability value using a standard F distribution. The column labelled t-stat reports the Engle and Granger cointegration test using critical values from MacKinnon (1996). † and ‡ indicate that the null of no cointegration is rejected at the 5 and 1%, respectively.

Table 4: Error correction models for deposit rates

	$\Delta d_{1,t}$	$\Delta d_{2,t}$	$\Delta d_{3,t}$	$\Delta d_{5,t}$	$\Delta d_{7,t}$	$\Delta d_{8,t}$	$\Delta d_{9,t}$	$\Delta d_{10,t}$	$\Delta d_{11,t}$	$\Delta d_{12,t}$	$\Delta d_{13,t}$	$\Delta d_{14,t}$	$\Delta d_{15,t}$
Const.	0.006	-0.005	-0.023	-0.023	-0.032	-0.008	-0.016	-0.017	0.009	-0.004	-0.009	-0.024	-0.029
	(0.020)	(0.019)	(0.014)	(0.018)	(0.019)	(0.019)	(0.021)	(0.012)	(0.027)	(0.013)	(0.018)	(0.015)	(0.018)
ect_{t-1}				-0.165				-0.065					-0.079
				(0.040)				(0.022)					(0.021)
ect^{+}_{t-1}	-0.143	-0.132	-0.035		-0.192	-0.154	-0.136		-0.317	-0.101	-0.128	-0.210	
	(0.042)	(0.048)	(0.023)		(0.065)	(0.047)	(0.051)		(0.059)	(0.030)	(0.048)	(0.068)	
ect^{-}_{t-1}	-0.026	-0.062	-0.131		-0.051	-0.062	-0.110		0.017	-0.036	-0.052	-0.094	
	(0.035)	(0.041)	(0.031)		(0.024)	(0.043)	(0.061)		(0.072)	(0.028)	(0.042)	(0.032)	
$\Delta d_{i,t-1}$	0.049	0.068	0.020	-0.183	-0.112	0.012	-0.037	0.256	0.043	0.231	-0.015	-0.083	0.008
	(0.083)	(0.083)	(0.081)	(0.083)	(0.082)	(0.083)	(0.083)	(0.085)	(0.085)	(0.083)	(0.083)	(0.082)	(0.083)
$\Delta d_{i,t-2}$	0.102	-0.008	0.096	0.039	-0.034	0.022	-0.028	0.085	-0.076	0.055	-0.166	0.062	-0.019
	(0.082)	(0.083)	(0.080)	(0.081)	(0.079)	(0.078)	(0.081)	(0.084)	(0.082)	(0.079)	(0.080)	(0.079)	(0.076)
Δib_{t}	0.392	0.431	0.353	0.554	0.223	0.358	0.419	0.419	0.640	0.284	0.270	0.347	0.229
	(0.069)	(0.070)	(0.060)	(0.080)	(0.084)	(0.072)	(0.073)	(0.056)	(0.121)	(0.048)	(0.070)	(0.064)	(0.075)
Δib_{t-1}	0.072	0.003	0.121	0.218	0.226	0.291	0.055	0.003	-0.054	0.048	0.301	0.175	0.337
	(0.077)	(0.080)	(0.068)	(0.092)	(0.092)	(0.077)	(0.082)	(0.065)	(0.136)	(0.051)	(0.072)	(0.072)	(0.082)
Δib_{t-2}	0.031	0.133	0.007	0.101	0.187	-0.038	0.104	0.027	0.458	0.003	0.014	0.032	0.091
	(0.075)	(0.077)	(0.067)	(0.089)	(0.092)	(0.079)	(0.082)	(0.065)	(0.124)	(0.051)	(0.074)	(0.072)	(0.086)
$\bar{R}^2$	0.420	0.426	0.449	0.443	0.304	0.456	0.413	0.593	0.413	0.505	0.378	0.420	0.404
LM(2)	1.833	1.148	0.198	1.274	1.027	2.998	0.351	0.106	0.141	0.138	5.524	2.713	0.665
p-value	[0.164]	[0.320]	[0.820]	[0.283]	[0.361]	[0.053]	[0.705]	[0.900]	[0.869]	[0.871]	[0.005]	[0.070]	[0.516]

Note: The error correction term is denoted *ect*. Standard errors are reported in parentheses. LM(2) is the Breusch-Godfrey LM test for residual serial correlation of up to order 2.

Table 5: Error correction models for lending rates

	Δl_2	Δl_5	Δl_8	Δl_{10}	Δl_{11}	Δl_{13}	Δl_{15}
Const.	0.107	-0.134	0.055	-0.087	-0.075	0.091	0.014
	(0.090)	(0.078)	(0.103)	(0.054)	(0.077)	(0.111)	(0.129)
ect^{+}_{t-1}	-0.159	-0.203	-0.334	-0.235	-0.144	-0.511	-0.367
	(0.044)	(0.076)	(0.098)	(0.061)	(0.043)	(0.111)	(0.082)
ect^{-}_{t-1}	0.042	-0.649	-0.047	0.283	0.020	-0.076	-0.058
	(0.048)	(0.152)	(0.054)	(0.085)	(0.045)	(0.095)	(0.060)
$\Delta l_{i,t-1}$	-0.387	-0.274	-0.171	-0.056	-0.209	-0.274	0.009
	(0.083)	(0.091)	(0.090)	(0.093)	(0.086)	(0.093)	(0.084)
$\Delta l_{i,t-2}$	0.036	-0.081	-0.142	-0.107	-0.061	-0.152	-0.062
	(0.086)	(0.080)	(0.084)	(0.096)	(0.094)	(0.082)	(0.083)
Δib_t	0.112	-0.173	0.520	0.374	0.085	-0.574	1.515
	(0.280)	(0.329)	(0.430)	(0.234)	(0.326)	(0.459)	(0.557)
Δib_{t-1}	0.117	0.630	-0.292	0.142	-0.226	0.790	-0.895
	(0.278)	(0.349)	(0.458)	(0.255)	(0.346)	(0.494)	(0.600)
Δib_{t-2}	0.595	0.197	-0.049	0.107	0.043	0.239	0.091
	(0.269)	(0.342)	(0.428)	(0.237)	(0.328)	(0.468)	(0.566)
$\bar{R}^2$	0.262	0.339	0.134	0.165	0.083	0.313	0.152
LM(2)	0.056	0.912	1.978	3.995	0.348	0.063	0.522
p-value	[0.945]	[0.404]	[0.143]	[0.021]	[0.707]	[0.939]	[0.594]

Note: The error correction term is denoted *ect*. Standard errors are reported in parentheses. LM(2) is the Breusch-Godfrey LM test for residual serial correlation of up to order 2.

Figure 1: Interest rate of usury and range of variation of deposit and lending rates