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Fisher's hypothesis, survey-based expectations, and asymmetric adjustments

Empirical evidence from South Africa

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Abstract: Our study re-examines Fisher’s hypothesis for South Africa in the post-inflation targeting era and presents two noteworthy empirical contributions. Firstly, we examine the Fisher effect by making use of survey-based inflation expectations data for financial analysts, the business sector, trade unions, and households. Secondly, we examine both short-run and long-run asymmetric cointegration effects in Fisher’s relation using the nonlinear autoregressive distributive lag model as an econometric framework. Our full quarterly sample (2002:Q1–2019:Q4) finds interest rates to respond more aggressively to falling expectations than rising one, with a full Fisher effect found for financial analysts, partial effects for households and the business sector, and no effects for trade unions. However, after splitting the data into pre- and post-financial crisis periods, we observe changing dynamics in which interest rates respond more aggressively to rising inflation, with partial effects also being found for trade unions. Policy recommendations are offered.

Key words: Fisher effect, survey-based inflation expectations, interest rates, nonlinear autoregressive distributive lag model, South Africa.

JEL classification: C12, C22, E52, E58.

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1 Introduction

The Fisher effect, which dates back to the independent work of Fisher (1930), is one of the most investigated research topics in monetary economics and is based on a simple mathematical premise linking interest rates to expected inflation. According to Fisher's hypothesis, over the long run, nominal interest rates are said to have a one-for-one relationship with expected inflation rates, implying that long-run real interest rates remain unchanged over their steady-state equilibrium. Consequentially, if this hypothesis is proven to be true, then nominal interest rates can be deemed an unbiased estimator for inflation and, in turn, interest rates can act as a hedge against inflation for financial market participants such as savers and investors.

Overall, the Fisher effect is important for the stability of financial markets and represents a key relationship within monetary policy frameworks. Central banks, which tend to be forward-looking, primarily base their policy decisions on future inflation forecasts. This is more prominent in inflation targeting (IT) economies, whose primary policy mandate is to stabilize inflation within a predetermined 'inflation band-width'. Under these frameworks, if inflation is forecast to breach the upper boundary of the inflation band-width, policy makers raise their interest rates, and once inflation expectations are curbed downwards towards their targeted values, then central banks can pursue expansionary policy by lowering policy rates. Fisher's hypothesis speculates that if central banks can ensure a 'one-for-one' co-movement between interest rates and expected inflation, then the purchasing power of savers and investors will be protected since interest earned from financial institutions would be high enough to offset any changes in the prices of goods and services.

In our study, we investigate the Fisher effect for the South African Reserve Bank (SARB), which is the oldest central bank in Africa and the continent's first reserve bank to adopt a fully-fledged IT framework as official monetary policy in 2002. Whilst we are aware of previous studies by Bahmani-Oskooee et al. (2016), Bayat et al. (2018), Mitchell-Innes et al. (2007), Phiri and Lusanga (2011), Wesso (2000), and Yaya (2015), which all empirically tested Fisher's hypothesis for South African data and found conflicting empirical evidence, the debate concerning the validity of the Fisher effect in South Africa is far from reaching a consensus and is still open to further deliberation. Our study addresses two fundamental issues which have been ignored by the previous South African literature.

Firstly, we address the issue of measures of inflation expectations used in estimating Fisher's equation, in relation to which we note that previous South African studies construct aggregated measures of inflation expectations extracted from actual inflation rates. However, in reality, different economic agents form various inflation expectations across different time horizons (Kabundi et al. 2015; Miyajima and Yetman 2019) and hence it is important to know whether interest rates respond differently to expectations formed by the different market participants at different forecast periods. The previous studies by Bond and Somlen (1992), Darin and Hetzel (1995), Gibson (1972), Kaliva (2008), Lai (1997), Peek and Wilcox (1983), and Soderlind (1998) used survey-based inflation forecasts to investigate the Fisher effect and found that survey-based data circumvents the problems of systematic forecasting errors produced from econometric-based forecasts. Our study contributes to this line of research by adopting a disaggregated approach to examining the Fisher effect, using survey-based inflation expectations collected by the Bureau of Economic Research (BER) at Stellenbosch University for four broad market participants in the South African economy, namely: i) financial analysts, ii) the business sector, iii) trade unions, and iv) household participants. We also address the issue of inflation expectation forecast horizons, in which our study differs from previous works by using multiple forecast horizons which more closely align with the actual forecast horizons used by the SARB in practice.

Secondly, our paper addresses the issue of possible nonlinear Fisher effects in the data as most of the previous South African literature assumes a monotonic response of interest rates to inflation expectations, which are typically captured using linear cointegration models such as the vector error correction model (VECM) framework of Johansen and Juselius (1990) (i.e. Mitchell-Innes et al. 2007; Wesso 2000) and the autoregressive distributive lag (ARDL) model of Pesaran et al. (2001) (i.e. Yaya 2015). We note exceptions for the studies of Phiri and Lusanga (2011), who apply the threshold vector error correction (TVEC) model of Hansen and Seo (2002), as well as the study by Bahmani-Oskooee et al. (2016), which applies an autoregressive distributive lag (ADL) test of Li and Lee (2010) for threshold cointegration effects. However, the TVEC estimation model only assumes asymmetries over short-run convergence processes whilst remaining linear in its long-run parameters, and the ADL test does not go beyond testing for significant threshold effects and, consequentially, does not produce any regression coefficients for empirical scrutiny.

Our study, circumvents empirical difficulties previously experienced in establishing long-run asymmetries in Fisher's relationship for South African data by relying on the nonlinear autoregressive model (NARDL) of Shin et al. (2014), which is nonlinear in both long-run and short-run parameter estimates and provides a more unified framework for examining asymmetric cointegration effects within time series data. Notably, this feature is in contrast to other competing nonlinear models used in the general literature to examine the nonlinear relationship between nominal interest rates and inflation expectations. For instance, the smooth transition regression (STR) frameworks of Choi and Saikkonen (2004), Fouquau et al. (2008), Gonzalez et al. (2005), Saikkonen and Choi (2004), and Terasvirta (1994), and previously used in the studies of Ahmad (2010), Christopoulos and Leon-Ledesma (2007), Kim et al. (2018), Nusair (2009), and an Yoon (2010), are limited to modelling asymmetries over the long run whilst ignoring short-run and equilibrium adjustment asymmetries. On the other hand, the momentum threshold autoregressive (MTAR) model of Enders and Granger (1998) and Enders and Siklos (2001) previously used by Bajo-Rubio et al. (2005) and Maki (2005), as well as the TVEC model of Balke and Fomby (1997) and Lo and Zivot (2001) used in the works of Dutt and Ghosh (2007) and Million (2004) assume short-run adjustment asymmetries whilst maintaining linearity over the long-run regression parameters.

Notably, the NARDL model further presents flexibility advantages over the preceding nonlinear frameworks as it is flexible enough to not require mutual integration of the time series and also produces robust estimates even with a relatively short sample of data. As far as we are concerned, only the recent studies by Ongan and Gocer (2018, 2019, 2020) have employed the NARDL model to estimate nonlinear Fisher effects for Canada, South Korea, and the USA, respectively. Our study builds on this recent work in the context of the South African economy and uses the NARDL model to investigate the Fisher effect between long-term government bond yield and disaggregated survey-based inflation expectations.

From a policy perspective, our strategy of investigating the nonlinear Fisher effect using the NARDL model applied to disaggregated survey-based data presents a more effective method of using Fisher's hypothesis to determine whether the SARB has been successful in anchoring inflation expectations of different economic agents. Firstly, as highlighted by Aron and Mellbauer (2007), Kabundi et al. (2015), Miyajima and Yetman (2019), and Reid (2009), expectations formed by different economic agents tend to be heterogenous, hence it is important for the Reserve Bank to understand how the expectations of different agents are shaped by monetary policy. Secondly, the use of the NARDL model more effectively models the different sources of nonlinearity between nominal interest rates and expected inflation identified in the literature. For instance, Coakley and Fuertes (2002) and Million (2004) attribute asymmetries in the Fisher effect and other monetary policy relationships to the inflation-targeting practices and opportunistic behaviour of central banks, whereby policy makers do not immediately manipulate interest rates to control inflation but, rather, wait for favourable external shocks before taking policy action. Moreover, Bec et al. (2002), Dolado et al. (2005), Kim et al. (2005), and Schaling (2004) developed models of nonlinear 'Taylor-type' policy rules in which interest rates adjust more than 'one-for-one' when inflation is expected to rise and respond in a less than 'one-for-one' fashion when expected inflation is falling. The NARDL model can more efficiently capture these described nonlinearities by partitioning the Fisher effect into two phases, one capturing the dynamics when inflation is rising and the other when inflation is failing. Overall, our empirical findings verify these nonlinear dynamics between nominal interest rates and inflation expectations, although we observe different asymmetric Fisher dynamics for different agent inflationary expectations and across different time periods.

We present the rest of our study as follows. The following section outlines the empirical framework used in our study. The third section of the paper presents the data and empirical findings. The study is concluded in the fourth section of the paper in the form of policy implications.

2 Empirical framework

Within his theory of variations in investments, Fisher (1930) advocated for the existence of the relationship between interest rates and the inflation rate. Under Fisher's hypothesis, nominal interest rates and the inflation rate are said to have a long-run 'one-for-one' relationship, with the real interest rate remaining constant. Two empirical approaches exist when testing for Fisher effects. Firstly, one can test the integration properties of the real interest rates and confirm Fisher effects if the series is found to be a mean-reverting I(0) process. However, the disadvantage of this approach is that it fails to validate whether nominal interest rates and inflation expectations do indeed move proportionately, as implied by the strict definition of the Fisher effect. Secondly, and in the approach adopted in our study, one can investigate the Fisher effect using a bivariate estimation regression between nominal interest rates (i_t) and inflation expectations (π_t^e) i.e.

$$i_t = \alpha + \beta\, \pi_t^e + \varepsilon_t \qquad\qquad \varepsilon_t \sim N(0, \sigma^2) \tag{1}$$

From equation (1), a full Fisher effect is said to exist if the regression coefficient satisfies the following condition $\beta_1 = 1$, whereas partial effects are assumed if $0 < \beta < 1$. These partial effects are attributed to the Mundell-Toni effect, under which the substitutability of bonds and real money balances results in a negative effect of anticipated inflation on real bond yields (i.e. wealth effects). It is also possible that $\beta_1 > 1$, and this more than 'one-for-one' relationship is attributed to the 'tax effects' described in Darby (1975) and Feldstein (1976), in which nominal interest rates rise more than proportionately to inflation to ensure that the after-tax real return remains unchanged. In our study, we assume an asymmetric response of nominal interest rates to inflation expectations, with the differing effects being dependent on whether inflation is increasing or decreasing. To do this, we follow Shin et al. (2014) and propose that inflation expectations, π_t^e , can be decomposed into partial sum processes of positive and negative changes (i.e. $i_t = \pi_0^e + \pi_t^{e+} + \pi_t^{e-}$), such that equation (1) can be re-specified as the following long-run asymmetric model:

$$i_t = \alpha_0 + \beta^+ \pi_t^{e+} + \beta^- \pi_t^{e-} + e_t \tag{2}$$

where π_t^{e+} and π_t^{e-} are partial sum processes of positive and negative changes in π_t^e defined as:

$$\pi_t^{e+} = \sum_{j=1}^{i} \Delta\, \pi_j^{e+} = \sum_{j=1}^{i} \max\,(\Delta\pi_j, 0) \tag{3}$$

$$\pi_t^{e-} = \sum_{j=1}^{i} \Delta \pi_j^{e-} = \sum_{j=1}^{i} \min(\Delta\pi_j^{e}, 0) \tag{4}$$

The NARDL (p, q) in-levels transformation of regression (4) can be given as:

$$i_t = \sum_{j=1}^{p} \phi_i \pi_{t-j} + \sum_{j=1}^{p} \left(\theta_j^{+} \pi_{t-j}^{e+} + \theta_j^{-} \pi_{t-j}^{e-}\right) + \zeta_t \tag{5}$$

where ϕ_i is the autoregressive parameter, θ_j^{-} and θ_j^{-} are the asymmetric distributive-lag parameters, and ζ_t is a well-behaved error with properties $N \sim (0, \sigma^2)$. From, equation (5), the unrestricted error correction representation can be expressed as:

$$i_t = \sum_{j=1}^{p} \rho_i i_{t-j} + \theta_j^{+} \pi_{t-j}^{e+} + \theta_j^{-} \pi_{t-j}^{e-} + \sum_{j=1}^{p-1} \lambda_i \Delta i_{t-j} + \sum_{j=0}^{q-1} (\alpha_j^{+} \Delta\pi_{t-j}^{e+} + \alpha_j^{-} \Delta\pi_{t-j}^{e-}) + \zeta_t \tag{6}$$

$$= \sum_{j=1}^{p} \rho_i \ \xi_{t-j} + \sum_{j=1}^{p-1} \lambda_i \Delta i_{t-j} + \sum_{j=0}^{q-1} (\alpha_j^{+} \Delta\pi_{t-j}^{e+} + \alpha_j^{-} \Delta\pi_{t-j}^{e-}) + \zeta_t \tag{7}$$

where $\xi_{t-j} = i_t - \theta_j^{+} \pi_{t-j}^{e+} - \theta_j^{-} \pi_{t-j}^{e-}$ is the asymmetric error correction term and the asymmetric long-run parameters are computed as $\beta^{+} = -(\theta^{+}/\rho)$ and $\beta^{-} = -(\theta^{-}/\rho)$. Note that the NARDL model admits three types of nonlinearity, namely: i) long-run or reaction asymmetry, ii) short-run asymmetry, and iii) adjustment asymmetry, and Shin et al. (2014) develop a battery of testing procedures to test the significance of the different forms of asymmetries. Firstly, the authors propose two tests for asymmetry cointegration in the NARDL model which are merely asymmetric extensions of the linear cointegration tests for linear ARDL models presented in Pesaran et al. (2001). On the one hand, there is the t-statistic, which tests the null hypothesis $\rho = 0$ against the alternative $\rho < 0$ and evaluates the significance of equilibrium adjustment asymmetries via the nonlinear ECM. On the other hand, there is the asymmetric version of bounds test for cointegration, which is a F-test of the joint null hypothesis, $\rho = \theta^{+} = \theta^{-}$. The test statistics which evaluate both sets of hypotheses are denoted t_{BDM} and F_{PSS}, respectively. Secondly, the authors propose two additional sets of tests which evaluate the long-run and short-run asymmetries. On the one hand, there is the Wald test, which evaluates the null hypotheses of long-run or reaction symmetry, and which imposes the restriction $\beta^{+} = \beta^{-} = \beta$, and reduces equations (6) and (7) to:

$$i_t = \sum_{j=1}^{p} \rho_i (i_{t-j} - \beta\pi_{t-j}^{e}) + \sum_{j=1}^{p-1} \lambda_i \Delta i_{t-j} + \sum_{j=0}^{q-1} (\alpha_j^{+} \Delta\pi_{t-j}^{e+} + \alpha_j^{-} \Delta\pi_{t-j}^{e-}) + \zeta_t \tag{8}$$

and the test statistic evaluating the null hypotheses $(H_{LR}^{S}: \beta^{+} = \beta^{-})$ is denoted as W_{LR}. On the other hand, the Wald test, which tests for short-run symmetry, imposes the restriction $\sum_{i=0}^{q-1} \alpha_j^{+} = \sum_{i=0}^{q-1} \alpha_j^{-} = \sum_{i=0}^{q-1} \alpha_j$ hence reducing equations (6) and (7) to:

$$i_t = \sum_{j=1}^{p} \rho_i i_{t-j} + \theta_j^{+} \pi_{t-j}^{e+} + \theta_j^{-} \pi_{t-j}^{e-} + \sum_{j=1}^{p-1} \lambda_i \Delta i_{t-j} + \sum_{j=0}^{q-1} (\alpha_j \Delta\pi_{t-j}^{e}) + \zeta_t \tag{9}$$

and the test statistic evaluating the null hypotheses $(H_{SR}^{S}: \beta^{+} = \beta^{-})$ is denoted as W_{SR}.

3 Empirical data and estimation results

3.1 Empirical data

As previously highlighted, we employ survey-based inflation expectations collected from the BER at Stellenbosch University to examine the Fisher effect for the South African economy. The BER started collecting surveyed data for inflation expectations in 2000 after being commissioned by the

SARB (Leshoro 2018). Notably, the actual face-to-face interviews are conducted by marketing research firm AC Nielsen (Du Plessis et al. 2018). The survey is conducted in quarterly time-frequencies for 2,500 households (i.e. *HH*), 34 financial analysts (i.e. *FA*), 480 business people (i.e. *BS*) and 37 trade unions representatives (i.e. *TU*), and the design of the surveys is similar to that of the Philadelphia Federal Bank Livingstone Survey and the University of Michigan survey of consumers (Miyajima and Yetman 2019; Reid et al. 2020). The forecasts are conducted for current, 12-month ahead, 24-month ahead, and, more recently, for 60-month ahead periods, and data is available in time series data format from the SARB online statistical database. Our study makes use of inflation expectations quarterly data spanning the period from 2002:Q1 to 2019:Q4 for all four surveyed market participants using current, 12-month ahead, and 24-month ahead forecasts. As a measure of nominal interest rates, used as the dependent variable in our Fisher equation, we employ the 10-year yield on the government bond (i.e. *BOND*) and, as argued by Ongan and Gocer (2020), this bond rate provides advantages such as being a proxy for other interest rates which dominate financial markets as well as reflecting investor confidence concerning their current and future expectations of the economy.

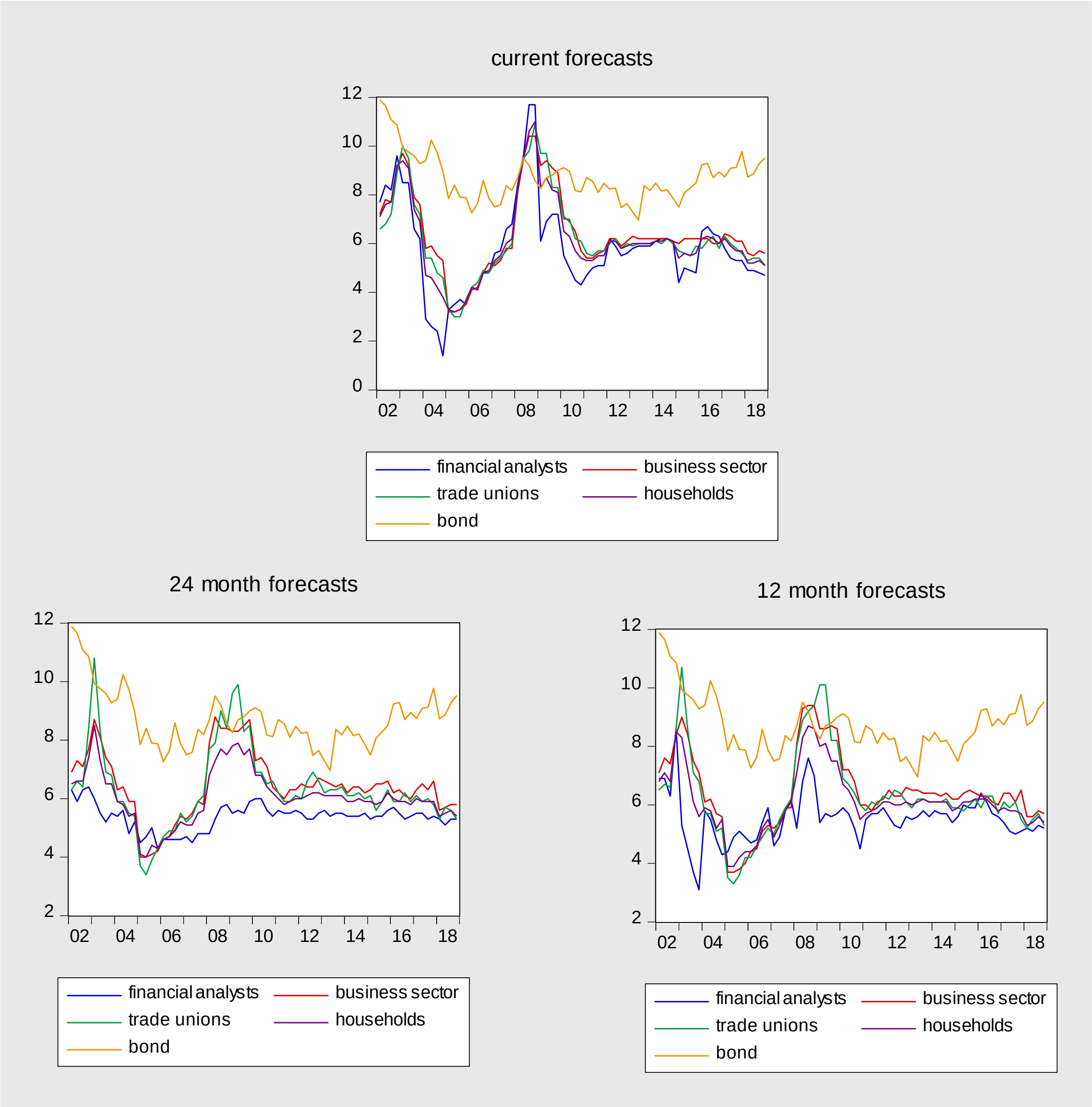
Table 1 presents the summary statistics of the time series along with their correlation coefficients with the 10-year yield on the government bond, and Figure 1 presents their corresponding time series plots. We find that since the adoption of the IT framework in 2002, the average of inflation expectations only lies within the SARB's set 3–6 per cent target range for financial analysts, whilst for business, trade unions, and households, average expectations breach the upper 6 per cent threshold of the target. As explained in Miyajima and Yetman (2019), the backward-looking formation of expectations by businesses, trade unions, and households are primarily determined by the growth in wage rates, which have generally remained higher than actual inflation. On the other hand, Kabundi et al. (2015) find that expectations for financial analysts are likely to lie within the target boundary since these experts are more informed on the operations of the Reserve Bank's IT policy. Based on the reported standard deviations, there is generally low volatility amongst expectations form the different agents, even though the ranking of volatility amongst the different agents varies with the forecast horizon. The reported correlation coefficients produce estimates ranging from 0.29 (trade union 24-month forecasts) to 57 (financial analyst 24-month forecasts), all of which imply a less than 'one-for-one' partial Fisher relationship between the variables, which corresponds to the previous findings from Mitchell-Innes et al. (2007). We, however, treat this evidence as being merely preliminary to our main analysis.

Table 1: Summary statistics and correlation of the variables

Time series	Mean	Median	Max.	Min.	Std. Dev.	Skew	Kurt	JB	Prob.	Correlation with BOND
BOND	8.72	8.58	11.89	6.96	0.97	1.06	4.60	2.06	0.46	1.00
FA	5.80	5.65	11.87	1.40	1.89	0.79	4.58	4.26	0.73	0.34
FA(+12)	5.52	5.60	8.60	3.10	0.82	0.59	6.09	3.30	0.13	0.34
FA(+24)	5.36	5.40	6.40	4.30	0.42	-0.26	3.17	0.87	0.65	0.57
BS	6.41	6.20	10.40	3.20	1.62	0.60	3.35	4.43	0.10	0.45
BS(+12)	6.48	6.30	9.40	3.70	1.30	0.31	3.24	1.28	0.53	0.43
BS(+24)	6.65	6.40	8.80	4.00	1.12	0.06	3.13	0.09	9.95	0.41
TU	6.24	6.00	10.90	3.00	1.65	1.65	3.75	2.65	0.61	0.36
TU(+12)	6.31	6.10	10.70	3.30	1.47	0.93	4.26	1.29	0.30	0.33
TU(+24)	6.33	6.10	10.80	3.40	1.34	0.95	4.75	1.02	0.19	0.29
HH	6.14	6.00	11.00	3.20	1.64	1.64	0.84	0.86	0.90	0.39
HH(+12)	6.11	6.00	8.70	3.90	1.05	0.57	3.61	4.75	0.90	0.42
HH(+24)	6.05	5.95	8.50	4.00	0.91	0.24	3.41	1.19	0.55	0.40

Source: authors' computation.

Figure 1: Inflation expectations (current, 12-month, and 24-month) and the bond rate



Source: authors' computation.

Whilst Pesaran et al. (2001) and Shin et al (2014) demonstrate that unit root tests are not required to validate the mutual integration of the time series a priori to estimating linear and nonlinear ARDL cointegration models, we, however, find it necessary to perform unit root tests on the first differences on the variables. This is just to ensure that none of the employed time series data violates the condition of being integrated of the order I(2) or higher. Table 2 presents the findings of the ADF, PP, and DF-GLS test performed on the first differences of the time series variables, which displays overwhelming evidence in favour of the series not being integrated of an order higher than I(1).

Table 2: Unit root test results on first differences of series

	ADF		PP		DF-GLS	
	Intercept	Intercept + trend	Intercept	Intercept + trend	Intercept	Intercept + trend
ΔBOND	-8.58***	-8.91***	-8.81***	-9.66***	-8.38***	-9.04***
ΔFA	-7.98***	-7.89***	-7.94***	-7.88***	-7.11***	-7.68***
ΔFA(+12)	-7.62***	-7.56***	-15.43***	-15.72***	-7.68***	-7.98***
ΔFA(+24)	-10.61***	-10.57***	-10.74***	-10.62***	-10.26***	-9.08***
ΔBS	-5.99***	-5.94***	-5.96***	-5.94***	-5.08***	-5.68***
ΔBS(+12)	-6.51***	6.46***	-6.53***	-6.48***	-5.49***	-6.17***
ΔBS(+24)	-7.46***	-7.40***	-7.47***	-7.41***	-6.46***	-7.17***
ΔTU	-3.38**	-3.35*	5.93***	-5.89***	-3.31***	-3.36**
ΔTU(+12)	-6.57***	-6.52***	-6.57***	-6.52***	-6.52***	-6.57***
ΔTU(+24)	-7.74***	-7.68***	-7.74***	-7.69***	-5.47***	-7.72***
ΔHH	-6.39***	-6.34***	-6.35***	-6.30***	-3.24***	-6.18***
ΔHH(+12)	-6.69***	-6.64***	-6.70***	-6.64***	-6.17***	-6.52***
ΔHH(+24)	-7.21***	-7.16***	-7.25***	-7.19***	-4.96***	-7.22***

Note: (*), (**), and (***) denotes significance at 10%, 5%, and 1% significance level, respectively.

Source: authors' computation.

3.2 Estimation results

Prior to presenting the main NARDL estimates, we provide baseline estimates from the linear ARDL model regressions of Pesaran et al. (2001) and report the findings in Table 3. Panel A presents the findings using inflation expectations from financial analysts, and we only observe a significant long-run 'one-for-one' relationship between interest rates and 24-month forecasts (equation 3), whilst short-run partial effects are found with 12-month (equation 2) and 24-month forecasts (equation 3). Panel B presents the results using inflation expectations from the business sector, which reveal partial long-run and short-run Fisher effects across current (equation 4), 12-month (equation 5), and 24-month forecasts (equation 6). The findings from the trade sector, reported in Panel C, reveal no significant short-run and long-run Fisher effects across all estimated regressions, whereas those from households, reported in Panel D, reveal partial long-run effects for 24-month (equation 12) ahead forecasts and partial short-run effects for 12-month (equation 11) ahead forecasts. For all estimated regressions, we also observe error correction term (ECT) estimates which lie between -0.20 and -0.27, hence indicating that the variables completely revert back to their steady-state equilibrium approximately four to five months after experiencing a shock to the series. Overall, our baseline analysis indicates that since the inception of the inflation targeting regime, the Reserve Bank has responded 'one-for-one' with financial analysts and less than 'one-for-one' with the business sector and households, and is irresponsive to trade unions. However, these preliminary estimates do not account for possible asymmetric relations whereby the responses of interest rates to expectations differ during periods of rising and falling expectations.

Table 3: Baseline ARDL estimates

Dependent variable: 10-year yield on government bond			
Panel A: Financial analysts	***(1) Current***	***(2) +12 months***	***(3) +24 months***
Long-run			
FA	0.14 (0.25)	0.30 (0.41)	1.01*** (0.00)
Short-run			
ΔFA	0.03 (0.31)	0.23*** (0.00)	0.27*** (0.02)
ECT(-1)	-0.23*** (0.00)	-0.20*** (0.00)	-0.27*** (0.00)
Panel B: Businesses	***(4) Current***	***(5) +12 months***	***(6) +24 months***
Long-run			
BS	0.24* (0.04)	0.33** (0.01)	0.38** (0.01)
Short-run			
ΔBS	0.06* (0.09)	0.09** (0.05)	0.09*** (0.03)
ECT(-1)	-0.26*** (0.00)	-0.26*** (0.00)	-0.26*** (0.00)
Panel C: Trade unions	***(7) Current***	***(8) +12 months***	***(9) +24 months***
Long-run			
TU	0.18 (0.16)	0.19 (0.19)	0.15 (0.34)
Short-run			
Δ TU	0.04 (0.23)	0.05 (0.27)	0.03 (0.40)
ECT(-1)	-0.24*** (0.00)	-0.23*** (0.00)	-0.23*** (0.00)
Panel D: All sectors	***(10) Current***	***(11) +12 months***	***(12) +24 months***
Long-run			
HH	0.19 (0.13)	0.27 (0.22)	0.37*** (0.00)
Short-run			
Δ HH	0.05 (0.20)	0.36*** (0.00)	0.09 (0.13)
ECT(-1)	-0.24*** (0.00)	-0.22*** (0.00)	-0.25*** (0.00)

Note: (*), (**), and (***) denotes the lower and upper bounds at 10%, 5% and 1% significance level, respectively. Parentheses () denotes the p-values.

Source: authors' computation.

Table 4 presents our main NARDL estimation results in which we distinguish Fisher effects between periods of rising expectations and falling expectations over different forecast horizons. From Panel A, which reports the findings for financial analysts, we observe significant nonlinear long-run Fisher effects for 24-month expectations (equation 3) in which interest rates respond more aggressively (i.e. 'greater-than-unitary' Fisher effect) during periods of falling expectations, whilst they respond 'one-for-one' during periods of falling expectations. However, over the short, run we note significant partial effects only for periods of rising expectations (i.e. FA^{POS}) for the 12-month forecast horizon (equation 2), whilst for 24-month forecasts (equation 3), we observe partial Fisher effects during both periods of rising (i.e. FA^{POS}) and falling (i.e. FA^{NEG}) expectations. From Panel B, we find significant partial long-run and short-run Fisher effects during both rising and falling expectations across all forecast horizons with the exception of current-period expectations over the short run (equation 4) whose coefficient estimates, ΔFA^{POS} and ΔFA^{NEG}, are statistically insignificant. In Panel C, we note insignificant short-run and long-run estimates for trade union expectations across all forecast horizons, whereas in Panel D expectations of households produce statistically significant partial long-run Fisher effects during periods of both rising and falling expectations over 12-month (equation 11) and 24-month (equation 12) forecast horizons, as well as for 12-month forecasts over both rising and falling periods. Collectively, our results are in line with those found for the linear ARDL estimates, in that we establish a full Fisher effect for expectations from financial analysts, partial effects for the business sector and household expectations, and no Fisher effect for trade union expectations. Moreover, our nonlinear estimates indicate a slightly more aggressive response of interest rates during periods of falling expectations in comparison to periods of rising expectations where the interest rates respond less aggressively to expectations.

Table 4: Non-linear ARDL estimates

Panel A: Financial analysts	***(1) Current***	***(2) +12 months***	***(3) +24 months***
Long-run			
FA^{POS}	0.15 (0.28)	0.12 (0.71)	0.99*** (0.00)
FA^{NEG}	0.14 (0.28)	0.06 (0.87)	1.08*** (0.00)
Short-run			
ΔFA^{POS}	0.03 (0.32)	0.34*** (0.01)	0.31*** (0.02)
ΔFA^{NEG}	0.03 (0.33)	0.01 (0.87)	0.33** (0.04)
ECT(-1)	-0.23*** (0.00)	-0.23*** (0.00)	-0.31*** (0.00)
Panel B: Businesses	***(4) Current***	***(5) +12 months***	***(6) +24 months***
Long-run			
BS^{POS}	0.23* (0.05)	0.32** (0.02)	0.37*** (0.01)
BS^{NEG}	0.24* (0.04)	0.33** (0.01)	0.38*** (0.01)
Short-run			
ΔBS^{POS}	0.06 (0.11)	0.08* (0.05)	0.09** (0.03)
ΔBS^{NEG}	0.06 (0.11)	0.09* (0.06)	0.09* (0.05)
ECT(-1)	-0.26*** (0.00)	-0.26*** (0.00)	-0.25*** (0.00)
Panel C: Trade unions	***(7) Current***	***(8) +12 months***	***(9) +24 months***
Long-run			
TU^{POS}	0.19 (0.17)	0.19 (0.20)	0.15 (0.35)
TU^{NEG}	0.18 (0.18)	0.19 (0.24)	0.15 (0.36)
Short-run			
ΔTU^{POS}	0.04 (0.23)	0.05 (0.25)	0.04 (0.40)
ΔTU^{NEG}	0.04 (0.25)	0.04 (0.27)	0.03 (0.42)
ECT(-1)	-0.24*** (0.00)	-0.23*** (0.00)	-0.22*** (0.00)
Panel D: All sectors	***(10) Current***	***(11) +12 months***	***(12) +24 months***
Long-run			
HH^{POS}	0.20 (0.14)	0.39*** (0.03)	0.35* (0.07)
HH^{NEG}	0.18 (0.14)	0.41*** (0.02)	0.37* (0.06)
Short-run			
ΔHH^{POS}	0.05 (0.19)	0.10* (0.06)	0.09 (0.14)
ΔHH^{NEG}	0.05 (0.21)	0.11* (0.08)	0.09 (0.16)
ECT(-1)	-0.24*** (0.00)	-0.27*** (0.00)	-0.25*** (0.00)

Note: (*), (**), and (***) denotes the lower and upper bounds at 10%, 5%, and 1% significance level, respectively. Parentheses () denotes the p-values.

Source: authors’ computation.

3.3 Sensitivity analysis

In this section of the paper, we present a sensitivity analysis in which we segregate our empirical data into two sub-samples corresponding to the pre-crisis and post-crisis periods. Table 5 presents the NARDL estimates based on the sub-samples.

The findings for financial analyst expectations reported in Panel A reveal switching dynamics in Fisher's relationship between the pre-crisis and post-crisis data. We particularly observe significant short-run and long-run Fisher effects during periods of falling expectations across all forecast horizons (equations 1–3), whereas for post-crisis data significant Fisher effects are established during periods of rising expectations (equations 4–6). We note a 'close-to-unity' value on the long-run coefficient for the 24-month forecast expectations during periods of falling expectations in the post-crisis period (equation 3), whereas for post-crisis data a 'one-for-one' relationship is found for 12-month expectations (equation 5) during periods of rising expectations and a 'greater-than-unit' value for 24-month forecasts (equation 6) during periods of both rising and falling expectations. The later results particularly highlight that during the post-crisis era, interest rates responded over-aggressively towards longer-term expectation horizons during periods of both rising and falling expectations.

From the results reported for the business sector in Panel B, we observe similar switching dynamics between the pre-crisis and post-crisis data. Firstly, we find that significant long-run partial Fisher effects exist across current and 12-month forecast horizons (equations 7 and 8) for periods of falling expectations during the pre-crisis period. Secondly, the post-crisis data reveals significant, long-run, less than 'one-for-one' coefficients for 12-month (equation 11) and 24-month (equation 12) forecast horizons, respectively, during periods of rising expectations. Thirdly, over the short run, we generally observe a significant partial Fisher effect in which interest rates respond more aggressively during periods of falling expectations (i.e. $BUS^{NEG} > BUS^{POS}$) for the pre-crisis data and respond more aggressively during periods of rising expectations (i.e. $BUS^{POS} > BUS^{NEG}$) for the post-crisis period. Collectively, and similarly to the results previously reported for the financial analysts, we observe that during the post-crisis period, interest rates responded more aggressively towards longer-term expectation horizons during periods of rising expectations and less aggressively during periods of falling expectations.

From the results based on trade union expectations reported in Panel C, we also observe discrepancies in the estimation results between the pre-crisis and post-crisis periods. We note that during the pre-crisis periods there are no statistically significant long-run Fisher effects across all forecast horizons (equations 13–15), whereas for the post-crisis data significant effects emerge exclusively for rising inflation expectations at all forecast horizons (equations 16–18). Moreover, the short-run estimates are similar to those obtained for the business sector, whereby significant partial Fisher effects exist across both sub-samples and yet interest rates respond more (less) aggressively during periods of falling (rising) expectations during the pre-crisis (post-crisis) era. Overall, an important revelation from our findings is that the short-run dynamics between interest rates and trade union expectations only translate into long-run Fisher effects exclusively during the post-crisis period, albeit that a partial-relationship is observed.

Lastly, the results from the expectations based on households reported in Panel D similarly reflect the common finding of switching Fisher dynamics between the pre-crisis and post-crisis data. We particularly observe that the results reported in Panel D 'more-or-less' reflect those previously found using data for business sector expectations. Firstly, similarly to the findings previously obtained from business sector expectations in the pre-crisis period, significant long-run partial Fisher effects are found exclusively for the periods of falling expectations of 12-month and 24-month forecast horizons (equations 20 and 21) and yet insignificantly for periods corresponding to rising expectations. Secondly, similarly to the results obtained from business sector expectations for the post-crisis period, we find 'less-than-unity' effects during periods of both rising and falling expectations, even though the latter produces lower coefficient estimates. Thirdly, similarly to the previous results based on expectations from financial analysts, the business sector, and trade unions over the short run, we observe partial Fisher effects in both the pre-crisis and the post-crisis periods, with interest rates responding more aggressively during periods of falling expectations during the pre-crisis period and responding more-aggressively during periods of rising expectations. Overall, our findings indicate significant long-run Fisher effects for all market participants except trade unions during the pre-crisis period, whilst during the post-crisis period the Fisher effects for trade unions turn significant along with the other market participants.

Table 5: Nonlinear ARDL estimates: pre- and post-crisis samples

		Pre-crisis			*Post-crisis*	
Panel A: Financial analysts	***(1) Current***	***(2) +12 months***	***(3) +24 months***	***(4) Current***	***(5) +12months***	***(6) +24months***
Long-run						
FA^{POS}	0.15 (0.13)	0.17 (0.58)	0.53 (0.78)	0.40* (0.09)	0.78* (0.05)	1.06*** (0.00)
FA^{NEG}	0.26*** (0.00)	0.49*** (0.03)	0.98* (0.06)	0.22 (0.35)	0.92 (0.50)	1.12*** (0.01)
Short-run						
ΔFA^{POS}	0.18 (0.12)	0.08 (0.51)	0.47 (0.53)	0.13** (0.05)	0.12*** (0.00)	0.62*** (0.03)
ΔFA^{NEG}	0.13* (0.06)	0.22** (0.01)	0.53** (0.02)	0.07 (0.29)	0.17 (0.29)	0.74*** (0.00)
ECT(-1)	-1.15*** (0.00)	-0.45** (0.02)	-0.22** (0.01)	-0.32** (0.02)	-0.18** (0.03)	-0.35** (0.01)
Panel B: Business sector	***(7) Current***	***(8) +12 months***	***(9) +24 months***	***(10) Current***	***(11) +12months***	***(12) +24months***
Long-run						
BS^{POS}	0.48 (0.19)	0.57 (0.11)	0.39* (0.04)	0.52 (0.13)	0.69* (0.08)	0.89* (0.05)
BS^{NEG}	0.46*** (0.00)	0.55*** (0.00)	0.66*** (0.00)	0.33 (0.19)	0.50 (0.14)	0.62* (0.09)
Short-run						
ΔBS^{POS}	0.27* (0.07)	0.54* (0.05)	0.23 (0.88)	0.27* (0.06)	0.39* (0.04)	0.48** (0.03)
ΔBS^{NEG}	0.42** (0.02)	0.61*** (0.01)	0.43*** (0.02)	0.10 (0.10)	0.16* (0.09)	0.23** (0.07)
ECT(-1)	-0.79** (0.02)	-1.20** (0.03)	-1.29*** (0.00)	-0.30** (0.01)	-0.33** (0.01)	-0.38*** (0.00)
Panel C: Trade unions	***(13) Current***	***(14) +12 months***	***(15) +24 months***	***(16) current***	***(17) +12months***	***(18) +24months***
Long-run						
TU^{POS}	0.87 (0.65)	0.41 (0.23)	0.45 (0.61)	0.69** (0.03)	0.75*** (0.03)	0.70*** (0.03)
TU^{NEG}	0.34 (0.82)	0.36 (0.79)	0.27 (0.87)	0.25 (0.16)	0.27 (0.13)	0.27 (0.14)
Short-run						
ΔTU^{POS}	0.11 (0.27)	0.26*** (0.01)	0.02** (0.01)	0.26** (0.01)	0.27** (0.01)	0.23** (0.01)
ΔTU^{NEG}	0.24 (0.61)	0.65* (0.09)	0.41* (0.03)	0.09 (0.10)	0.10* (0.08)	0.09* (0.09)
ECT(-1)	-0.26* (0.08)	-0.75* (0.04)	-0.17** (0.02)	-0.37*** (0.00)	-0.37** (0.02)	-0.34** (0.01)

Panel D: All sectors	*(19) Current*	*(20) +12 months*	*(21) +24 months*	*(22) current*	*(23) +12months*	*(24) +24months*
Long-run						
HH^{POS}	0.68	0.01	0.27	0.83**	0.76	0.81***
	(0.77)	(0.98)	(0.56)	(0.04)	(0.11)	(0.01)
HH^{NEG}	-2.04	0.52***	0.85***	0.34	0.84	0.74***
	(0.69)	(0.00)	(0.02)	(0.12)	(0.15)	(0.03)
Short-run						
ΔHH^{POS}	0.09*	0.03	0.50	0.30***	0.63**	0.88**
	(0.07)	(0.98)	(0.31)	(0.02)	(0.03)	(0.01)
ΔHH^{NEG}	0.39*	0.36***	0.65*	0.12**	0.25*	0.32**
	(0.08)	(0.03)	(0.09)	(0.07)	(0.05)	(0.02)
ECT(-1)	-0.13**	-0.69***	-1.20**	-0.30***	-0.37***	-0.40***
	(0.05)	(0.00)	(0.05)	(0.00)	(0.01)	(0.00)

Note: (*), (**), and (***) denotes the lower and upper bounds at 10%, 5%, and 1% significance level, respectively. Parentheses () denotes the p-values.

Source: authors' computation.

3.4 Diagnostic tests

In this section of the paper, we present the findings of the diagnostic tests performed on our empirical regressions. For the linear ARDL estimates, we perform the F-test for Bounds cointegration effects described in Pesaran et al. (2001), the Jarque-Brea (J-B) test for normality, the Breusch-Godfrey test (B-G) for serial correlation, the autoregressive conditional heteroscedasticity test (ARCH) for heteroscedasticity, and Ramsey's reset test for correct functional form. The findings obtained from the diagnostic tests performed for the linear ARDL regressions are reported in Panel A of Table 6 and indicate significant ARDL cointegration effects, correct functional as well as normally distributed form, no existing autocorrelation, and homoscedastic residual terms in each of the estimated regressions. We attribute the later findings to the use of heteroscedastic-consistent standard errors in estimating our regression.

For the NARDL estimates, we perform similar diagnostic tests to those performed for the linear ARDL regressions, except that we perform the F-test for bounds nonlinear cointegration effects described in Shin et al. (2014), the tests for long-run asymmetric effects (W_{LR}) and the tests for short-run asymmetric effects (W_{SR}). The empirical findings of the diagnostic tests for the NARDL regression estimates are reported in Panel B for the full sample, Panel C for the pre-crisis period, and Panel D for the post-crisis period. Similarly to the results obtained for the linear ARDL estimates, we find that residuals from the nonlinear estimated regressions are normally distributed, as well as no serial correlation and homoscedastic error terms in all estimated regressions. Moreover, we observe significant nonlinear bounds cointegration effects as well as significant short-run and long-run asymmetries. Notably, these later results are an improvement over those previously obtained in Bahmani-Oskooee et al.'s (2016) significant nonlinear cointegration effects for South African data.

All in all, the diagnostic tests imply that we can interpret the obtained empirical results with a fair amount of confidence and provide associated policy implications based on the regression estimates.

Table 6: Diagnostic tests for non-linear ARDL estimates

Dependent variable: 10-year yield on government bonds												
Dependent variable	FA	FA (+12)	FA (+24)	BS	BS (+12)	BS (+24)	TU	TU (+12)	TU (+24)	HH	HH (+12)	HH (+24)
Equation	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Panel A: Linear ARDL estimates (full sample)												
F_{PSS}	6.43	5.52	6.36	6.62	6.56	6.69	6.67	6.74	6.71	6.55	6.32	6.74
JB test	0.76 (0.68)	0.15 (0.93)	1.31 (0.52)	1.22 (0.54)	1.24 (0.54)	1.31 (0.52)	1.10 (0.57)	1.08 (0.58)	0.97 (0.62)	1.00 (0.61)	0.97 (0.62)	1.20 (0.55)
LM test	0.68 (0.51)	0.68 (0.51)	0.26 (0.77)	0.722 (0.48)	0.62 (0.55)	0.63 (0.54)	0.74 (0.48)	0.72 (0.49)	0.71 (0.50)	0.72 (0.49)	0.57 (0.58)	0.68 (0.51)
ARCH test	0.81 (0.37)	0.29 (0.54)	0.49 (0.26)	0.53 (0.47)	0.46 (0.50)	0.58 (0.45)	0.83 (0.37)	0.89 (0.35)	0.92 (0.34)	0.69 (0.41)	1.22 (0.28)	0.77 (0.38)
Ramsey RESET Test	1.01 (0.32)	0.24 (0.63)	1.36 (0.25)	1.18 (0.24)	1.49 (0.22)	1.76 (0.19)	1.42 (0.24)	1.38 (0.25)	1.43 (0.24)	1.81 (0.28)	0.00 (0.97)	1.46 (0.23)
PANEL B: Nonlinear ARDL models (full sample)												
F_{PSS}	4.65	7.08	4.56	4.84	13.28	4.90	4.90	4.14	4.93	4.76	10.93	4.92
W_{LR}	4.85** (0.04)	7.24** (0.04)	6.45** (0.06)	8.74*** (0.00)	12.68*** (0.00)	5.87** (0.04)	13.41*** (0.02)	23.56*** (0.00)	18.28*** (0.00)	7.92*** (0.01)	6.04*** (0.02)	24.38*** (0.00)
W_{SR}	3.18** (0.05)	3.24* (0.08)	13.68*** (0.00)	18.52*** (0.01)	6.94*** (0.01)	6.02*** (0.02)	5.83** (0.06)	6.89*** (0.03)	6.03*** (0.02)	3.06* (0.08)	9.36*** (0.00)	4.55** (0.04)
JB test	0.74 (0.69)	0.34 (0.84)	1.54 (0.46)	1.26 (0.53)	1.27 (0.52)	1.32 (0.51)	1.08 (0.58)	1.07 (0.59)	0.95 (0.62)	0.98 (0.61)	1.23 (0.54)	1.25 (0.53)
LM test	0.83 (0.44)	0.88 (0.42)	0.13 (0.87)	0.72 (0.59)	0.61 (0.55)	0.65 (0.53)	0.85 (0.43)	0.85 (0.43)	0.84 (0.44)	0.81 (0.45)	0.48 (0.62)	0.65 (0.52)
ARCH test	0.75 (0.39)	0.09 (0.77)	0.51 (0.48)	0.55 (0.46)	0.47 (0.49)	0.58 (0.45)	0.82 (0.37)	0.88 (0.35)	0.89 (0.35)	0.67 (0.42)	0.46 (0.49)	0.81 (0.37)
Ramsey RESET Test	1.74 (0.24)	0.00 (0.97)	0.88 (0.36)	1.48 (0.23)	1.66 (0.21)	2.08 (0.16)	1.85 (0.18)	1.89 (0.18)	1.92 (0.17)	1.54 (0.22)	1.08 (0.30)	1.55 (0.22)
PANEL C: Nonlinear ARDL models (pre-crisis)												
F_{PSS}	7.86	14.65	12.45	4.17	7.36	6.10	4.53	8.66	22.62	7.64	35.45	3.93
W_{LR}	3.89* (0.08)	6.72** (0.05)	6.08** (0.04)	11.93*** (0.00)	7.11** (0.06)	23.84*** (0.00)	6.31** (0.05)	9.08*** (0.02)	8.15*** (0.03)	15.73*** (0.00)	3.11* (0.09)	7.81*** (0.02)
W_{SR}	5.44* (0.06)	6.07* (0.06)	9.49*** (0.02)	8.41** (0.00)	5.64* (0.09)	19.72*** (0.00)	12.25*** (0.00)	5.52* (0.08)	5.01* (0.08)	21.52*** (0.00)	6.29** (0.05)	7.28*** (0.01)
JB test	0.32 (0.86)	0.37 (0.73)	1.37 (0.50)	0.31 (0.86)	0.44 (0.80)	0.24 (0.88)	0.14 (0.93)	1.25 (0.54)	1.23 (0.54)	0.68 (0.71)	0.61 (0.73)	0.33 (0.85)
LM test	5.04 (0.45)	0.14 (0.87)	1.63 (0.28)	2.79 (0.13)	0.93 (0.82)	6.31 (0.12)	0.69 (0.56)	0.45 (0.12)	4.79 (0.30)	12.98 (0.19)	0.52 (0.72)	19.59 (0.16)
ARCH test	0.31 (0.58)	0.91 (0.35)	1.83 (0.19)	0.01 (0.91)	0.67 (0.42)	0.00 (0.98)	1.34 (0.27)	1.71 (0.21)	0.40 (0.53)	0.43 (0.52)	1.22 (0.28)	0.13 (0.72)
Ramsey RESET Test	0.11 (0.75)	1.95 (0.16)	1.46 (0.28)	0.16 (0.70)	0.00 (0.94)	0.00 (0.99)	0.88 (0.40)	2.19 (0.20)	3.45 (0.20)	0.32 (0.62	1.89 (0.19)	0.03 (0.88)
PANEL D: Nonlinear ARDL models (post-crisis)												
F_{PSS}	5.81	4.83	12.63	11.75	5.18	21.69	3.57	3.93	3.71	4.73	7.12	4.19
W_{LR}	10.53*** (0.00)	6.38*** (0.02)	26.44*** (0.00)	3.62** (0.07)	16.74*** (0.00)	8.14*** (0.02)	9.91*** (0.00)	4.07** (0.05)	14.85*** (0.00)	3.98** (0.05)	9.33*** (0.00)	4.85** (0.05)
W_{SR}	4.34** (0.04)	9.18*** (0.00)	17.74*** (0.00)	3.74** (0.06)	9.43*** (0.00)	5.98** (0.05)	7.51*** (0.02)	6.12** (0.05)	9.37*** (0.00)	6.97*** (0.03)	5.14** (0.06)	6.13*** (0.03)
JB test	0.77 (0.67)	0.01 (0.99)	0.58 (0.74)	0.47 (0.79)	0.41 (0.81)	0.33 (0.85)	0.66 (0.72)	0.45 (0.79)	0.48 (0.79)	0.58 (0.75)	0.40 (0.82)	0.38 (0.83)
LM test	0.09 (0.91)	0.58 (0.56)	0.69 (0.51)	0.49 (0.61)	0.26 (0.77)	0.29 (0.74)	0.28 (0.75)	0.39 (0.68)	0.39 (0.68)	0.11 (0.89)	0.43 (0.65)	0.29 (0.75)
ARCH test	0.03 (0.87)	0.67 (0.42)	0.11 (0.75	0.04 (0.84)	0.12 (0.73)	0.07 (0.79)	0.31 (0.86)	0.01 (0.92)	0.00 (0.95)	0.08 (0.78)	0.05 (0.82)	0.19 (0.66)
Ramsey RESET Test	0.05 (0.83)	0.34 (0.57)	0.25 (0.62)	0.07 (0.78)	0.12 (0.72)	0.37 (0.55)	0.04 (0.85)	0.02 (0.88)	0.00 (0.97)	0.00 (0.97)	0.07 (0.80)	0.28 (0.60)

Note: parentheses () states the p-value, S denotes stable, NS denotes not stable, JB denotes Jarque-Bera test, and LM denotes Lagrange Multiplier test.

Source: authors' computation.

4 Conclusions

Since the adoption of the IT framework by SARB in the early 2000s, the BER has been contracted to assemble survey-based inflation expectations data based on interviews conducted with financial analysts, business, trade unions, and households. In our study, we investigated for Fisher effects between the 10-year government bond rate and these surveyed measures of inflation expectations or current 1-year and 2-year forecast periods. Using quarterly data sourced over the period from 2002:Q1 to 2019:Q4, our study reveals that under the IT regime, nominal interest rates respond differently to the various economic agents, across different forecast horizons and across different times.

Firstly, when estimating the Fisher effect for the different agents using the linear ARDL model, we observe a full Fisher effect for financial analysts, partial effects for the business sector and households, and no effects for trade unions. Notably, all observed significant Fisher effects are found to be more prominent over longer forecast horizons i.e. 24-month horizon. Secondly, when we conduct these estimations using the more sophisticated nonlinear ARDL model, we find that the results from the linear ARDL model are replicated, except that now we further observe that interest rates respond more aggressively to increasing expectations of the economic agents and less so towards their falling expectations. Lastly, when segregating our empirical data into two sub-samples corresponding to the pre- and post- financial crisis periods, we observe discrepancies over our empirical estimates. On one hand, nominal interest rates are found to be more responsive to the falling inflation expectations across financial analysts, businesses, and households during the pre-crisis periods whilst being irresponsive to long-run expectations of trade unions. On the other hand, nominal interest rates are found to be more responsive to rising expectations of all economic agents in the post-crisis period.

Altogether, our findings have some important implications for monetary policy. For instance, our findings reveal how, throughout the implementation of the IT regime, the Reserve Bank has successfully anchored the expectations for financial analysts more than other economic agents. These findings have been supported by the recent studies by Kabundi et al. (2015) and Miyajima and Yetman (2019), who attribute this finding to the forward-looking nature of financial analysts in forming their expectations, whereas the other market participants are backward-looking in their formation of inflation expectations. Moreover, our findings also reveal how the dynamics of the Reserve Bank in responding to the inflation expectations of economic agents have changed for periods subsequent to the global financial crisis. Our findings indicate that, before the crisis, monetary authorities were more attentive to the falling expectations of financial analysts, business, and households whilst paying little attention to trade unions. However, in the post-crisis period, the SARB has been increasingly focusing on the rising expectations of economic agents including trade unions, even though a full Fisher effect is exclusively found only for financial analysts.

Based on these policy implications, we therefore recommend that the Reserve Bank focuses more on the behavioural aspects of anchoring economic agents. In particular, we recommend that future research and policy implementation should be directed towards the use of 'behavioural nudges' in curbing the expectations of price-making economic agents (businesses and trade unions). Recently, Reid et al. (2020) proposed the use of 'nudges' in the design of household expectation surveys. Future academic endeavours could extend the use of similar nudging mechanisms to the survey designs of businesses and trade unions. Moreover, future research and policy design could focus on how media outlets, such as newspapers, magazines, social media, television, and news broadcasts, can be more effectively used to curb the inflation expectations of different economics agents, particularly those market participants whose expectations are formed backwardly.

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