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Estimating and Forecasting Default Risk: Evidence from Jamaica

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Application of the Government of Jamaica Zero-coupon Curve to Modeling Yield Curve Risk

Oma Coke

Abstract

This study uses the Svensson (1994) method to estimate quarterly Government of Jamaica (GOJ) zero-coupon yield curves from March 2014 to December 2016. The Svensson (1994) method of estimation was used to obtain the parsimonious yield curve. The estimated spot rate curve is then incorporated into an interest rates stress testing framework to assess the impact on portfolio holdings of parallel and nonparallel shifts of the yield curve. The results of the stress testing exercise show that exposure to parallel shifts of the curve was higher across the respective market participant groups relative to nonparallel shifts. Additionally, deposit-taking institutions and securities dealers were more vulnerable to shifts in medium-term segment of the yield curve. The life insurance subsector was more vulnerable to the long end of the yield curve while the general insurance subsector exposures were equally weighted across the short to medium term segment of the curve.

Keywords: yield curve, key rate duration, financial stability.

JEL classification: F31, F32, F41.

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1. INTRODUCTION

The yield curve depicts the relation between bond yields against their maturity. It can be used as a benchmark for pricing bonds and in value analysis more generally. In practice, the estimation of a yield curve is often derived from observations of market prices in the government debt market. The use of the government's debt portfolio may be attributable to the fact that in most jurisdictions the government is the largest issuer of bonds; coupled with the perceived risk profile–theoretically risk-free and practically low risk. The yield curve is also a useful indicator for central banks as they are able to capture changes in market expectations of macroeconomic conditions, monetary policy and investors risk preferences.

In light of the aforementioned, this study addressed two objectives. Firstly, a yield curve for the period 2014Q1 to 2016Q4 is estimated using Government of Jamaica (GOJ) domestic issued Jamaican dollar (JMD) denominated bonds. To accomplish this objective, the study used the Svensson (1994) parametric model to infer GOJ's yield curve from domestic bond prices. The choice of Svensson model was motivated by the increased flexibility of the curve while maintaining the parametric properties of the curve that provides sound economic intuition. The estimation of GOJ yield curve is motivated by Kladivko (2010) who uses the Nelson-Siegel model for Czech Treasury yield curve from 1999 to the present and Gürkaynak et al. (2006) who use the Svensson model to estimate the US Treasury curve from 1961 to the present. Further motivation for this paper was garnered from Langrin (2007) who estimated multifactor versions of the Vašíček (1977) and the Cox, Ingersoll, and Ross (Cox et al., 1985) models of the term structure of interest rates for GOJ zero-coupon bond prices. The estimation by Langrin (2007) was conducted via state space modeling on daily GOJ domestic bond yields from September 24, 2004, to July 28, 2006, obtained from Bloomberg. Unlike Langrin (2007), which relies on an affine diffusion term structure modeling, this study relies on a cross-sectional approach to estimate the GOJ domestic zero-coupon yield curve.

Secondly, since interest rate risk can be captured by changes in the yield curve, this study considers estimation of the key rate durations of the GOJ's domestic bond portfolio. The study further assesses the impact of shifts in the yield curve guided by the key rate duration model on portfolio holding of domestic issues by market participant groups.

This approach adds to the existing work of Tracey (2009) who employs principal component analysis and key rate durations for assessing interest rate risk of holdings of both local and global GOJ bonds by Jamaica's banking sector.

This study is organized as follows. Section 2 reviews the fundamental concepts of the yield curve; Section 3 presents the Svensson modeling framework; Section 4 provides an overview of the data used in model including a detailed discussion of inherent issues; Section 5 presents the results of the estimation, including an assessment of the fit of the curve; Section 6 demonstrates the application of the key rate duration model in assessing the impact of yield curve shifts on portfolio holdings of JMD denominated domestic government issues for existing market participant groups in Jamaica's financial system; and Section 7 concludes.

2. YIELD CURVE BASICS

This section provides a review of the fundamental concepts of bond pricing and the development of a yield curve.

2.1 The Discount Function and Zero-coupon Yields

The pricing of a bond is conditional on the present value of its future cash flows. The interest rate or discount function used to calculate the present value depends on the yield offered on comparable securities in the market. The discount function is used to maintain real value across the time, that is, time value of money. In theory, the application of the discount function to value a zero-coupon bond that pays \$1 in n years can be written as:

$$P_t = \delta_t(n) = e^{-r_t(n)\times n}, \tag{1}$$

where $\delta_t(n)$ denotes the continuous discount function as at time t and $r_t(n)$ is the continuously compounded rate of return (yield) demanded by the investor for holding such investment until n periods ahead of time t (n denotes the time to maturity). The subscript t denotes the variability of the discount function. From Equation 1 above, one may apply the necessary transposition to get an expression for the continuously compounded yield (spot rate) on the zero-coupon bond:

$$r_t(n) = \frac{-ln\left[\delta_t(n)\right]}{n}. \tag{2}$$

In applying the concept of compounding to bond pricing, one may consider expressing yields on a coupon-equivalent basis. In this case, the compounding may be assumed to be m times per year instead of being continuous (for example, semiannual compounding implies that $m=2$, the payment of coupon is two times per year). Thus, we express the relation between the continuously compounded yield and the m-compounded coupon-equivalent as

$$r_t(n) = m \times \ln\left[1 + \frac{r_t^{ce}(n)}{m}\right], \tag{3}$$

where $\frac{r_t^{ce}(n)}{m}$ denotes the coupon-equivalent yield compounded m times per year. Similarly, the discount function is expressed as

4 $$\delta_t(n) = \frac{1}{\left[1 + \frac{r_t^{ce}(n)}{m}\right]^{m \times n}}.$$

Thus, the relation between yields and coupon-equivalent yields creates ease of mobility between continuously compounding and its coupon equivalent counterparts. The relation between yields and maturities are captured by the yield curve.

2.2 Coupon Bond and the Par Yield Curve

Similar to zero-coupon bonds, the pricing of a coupon-bearing bond is conditional on the discount function; thus, the price is the sum of the discounted future cash flows of the bond. For illustration, consider the price of a coupon-bearing bond with a nominal value of 100 and coupon payment of \$C $\left(C = \frac{100c}{m}\right)$ that matures in exactly n years from time t as follows:

5 $$P_t(n) = \sum_{i=1}^{m \times n} C\delta_t(i/m) + 100\delta_t(n),$$

where $\delta_t(i)$, $i = 1, 2, ..., n$, are discount functions for their respective maturities. Note that the yield on a coupon-bearing bond is dependent on the coupon rate that is assumed. One implication of this condition, as pointed out by Gürkaynak et al. (2006), is the disparity in the yields of bonds with identical maturities but varying coupon values.

The yields on a coupon-bearing bond can be expressed in terms of par yields. A par yield may be defined as the coupon rate at which a bond with a specific maturity would be traded at par; that is, the rate at which the present value of the bond is equivalent to its nominal value. Hence, given a coupon-bearing bond with a nominal value of \$100 and maturity n, the par yield is obtain as follows:

$$100 = \frac{100c_t(n)}{m}\sum_{i=1}^{m\times n}\delta_t(i/m) + 100\delta_t(n), \tag{6}$$

where $c_t(n)$ denotes the n year par yield. From the Equation 6, the par yield can be expressed as

$$c_t(n) = \frac{m[1-\delta_t(n)]}{\sum_{i=1}^{m\times n}\delta_t(i/m)}. \tag{7}$$

The par yield serves as a proxy for the quotation of yield on a coupon-bearing bond by financial market participants (Gürkaynak et al., 2006). As discussed, the yield curve, once estimated, may be presented as a zero-coupon yield curve or a par yield curve. The curvature of the yield curve will reflect the sensitivity of bond prices to interest rates and is measured by the bonds duration and convexity.

2.3 Duration and Convexity

The duration of a bond is a measure of the sensitivity of a bond's value to changes in interest rates. This measure, modified duration, can easily be derived from the Macaulay duration methodology. Frederick Macaulay (1938) defines duration (coined as the Macaulay duration) on coupon-bearing bond as the weighted average of the time (in years) that the investor must wait to receive their cash flows, that can be expressed as

$$D = \frac{1}{P_t(n)}\left[\sum_{i=1}^{m\times n}\frac{i}{m}\frac{c}{m}\delta_t(i/m) + n\delta_t(n)\right], \tag{8}$$

where $\frac{c}{m}$ denotes the annual coupon payment compounded m times per year for a bond instrument. Bonds that pay coupon has a duration that is less than its maturity while for the case of a zero-coupon bond, its duration is equal to its maturity. From Equation 8 it is observed that for constant maturity and spot rate, the modified duration is inversely related to the coupon rate, that is, higher coupon rate results in shorter duration for a given maturity. In the context of the application, the modified duration is mostly considered. Unlike the Macaulay duration, the modified duration primarily assumes that the expected cash flow of the bond does not change when the yield changes.

The modified duration can be defined in terms of the Macaulay duration as the duration of the bond divided by one plus the yield on the bond (for a selected compounded period):

9
$$D^M = \frac{D}{\left(1 + \frac{r_t^{ce}}{m}\right)}.$$

Duration in general captures a linear relation between price changes and yield change. Thus, the measure is accurate for changes in bond price relative to small changes in yield. The nonlinearity of the relation between bond prices and yield to maturity impedes on the accuracy of the duration measure to capture effective price changes relative to large changes in yield. The nonlinear relation between price and yield to maturity is effectively accounted for in the measure of convexity. So, in a simplistic point of view, convexity is used to measure the portion of the bond price change relative to the change in the yield to maturity that is not accounted for in the duration measure. This can be depicted through the second-order Taylor approximation of bond price changes with respect to yield:

10
$$\frac{\Delta P_t(n)}{P_t(n)} \approx -D^M \Delta y_t + \frac{1}{2} C \left(\Delta y_t\right)^2,$$

where $C = \frac{1}{P_t(n)} \frac{d^2 P_t(n)}{dy_t^2}$ is the convexity of the bond. Convexity accounts for the uncertainty in yields observed at the long end of the yield curve which results in the yield curve depicting a concave shape. An implication of this is that the capital gain from a decline in the yield is higher than the capital loss from an increase in the yield. Notably, bonds with longer maturity portraying higher convexity results at times in what is referred to as convexity bias. The greater the convexity bias is, the more concave the yield curve will become. More details of the impact of convexity on the functional form of the yield curve are provided below.

3. MODEL SELECTION AND OVERVIEW

The modeling of a yield curve can be broadly categorized into two groups: *1)* parsimonious models and *2)* spline-based models (see Waggoner, 1994). Between the two groups, one has to decide on their preference in regard to the trade-off between accuracy, which is an advantage of the latter, and smoothness, which is an advantage of the prior.

The Bank for International Settlements (BIS, 2005) reports that 9 out of 13 central banks which report their yield curve estimates to the BIS use the parsimonious approach. The popularity of parsimonious models among central banks may be attributed to the inherent property of the parsimonious approach in providing sufficiently smooth yield curves which are consistent with underlying macroeconomic conditions and investors' preferences. Spline-based methods on the other hand provide a richer precision in the fitting of the curve and is a preferred choice if one is interested in small pricing anomalies. However, spline-based yield curves may not be smooth enough and may oscillate considerably over daily intervals (Kladivko, 2010).

In this paper, the parsimonious approach to estimating the yield curve for Jamaica was adopted. Under this framework, the Nelson-Siegel (Nelson and Siegel, 1987) and Svensson (Svensson, 1994) models are presented throughout the remainder of this section.

In their seminal work on yield curves, Nelson and Siegel (1987) assumed that the functional form for the instantaneous forward rate is the solution of a second-order differential equation whose roots are equal:

$$f(\tau) = \beta_0 + \beta_1 e^{-\lambda\tau} + \beta_2 \lambda\tau e^{-\lambda\tau}, \quad (11)$$

where $f(\tau)$ is the instantaneous forward rate for the τ periods ahead, $\theta = (\beta_0, \beta_1, \beta_2, \lambda)$ is a vector of parameters to be estimated. Equation 11 may be classified as a three-component exponential function. The first component, β_0, is known as the level and may be defined as the limit of the forward rate as τ tends to infinity (that is, the asymptotic rate at which the forward rate and spot rate converges). The second component, $\beta_1 e^{-\lambda\tau}$, controls the slope of the forward rate curve and is a monotonically decreasing term (if β_1 is positive) or increasing term (if β_1 is negative). The third component, $\beta_2 \lambda\tau e^{-\lambda\tau}$, controls the location and size of the hump in the forward rate curve (β_2 determines the magnitude and sign of the hump while λ determines the location of the hump).

Integrating Equation 11 (with respect to τ) from 0 to τ and dividing the outcome by τ we get the continuously compounded spot rate curve:

$$i_c(\tau) = \beta_0 + \beta_1 \left(\frac{1 - e^{-\lambda\tau}}{\lambda\tau} \right) + \beta_2 \left(\frac{1 - e^{-\lambda\tau}}{\lambda\tau} - e^{-\lambda\tau} \right), \quad (12)$$

where the subscript c denotes continuity. From Equation 12, one can compute the corresponding discount function by applying the established relation:

$$\delta(\tau) = e^{-i_c(\tau)\tau}. \tag{13}$$

The discount function can be used to price outstanding issue with specific coupon rate and maturity dates. The asymptotic properties of the model provide rich economic intuition. The curve (forward or spot) by definition converges to finite limits from both ends. Note that:

$$\lim_{\tau\to\infty} f(\tau) \equiv \lim_{\tau\to\infty} i_c(\tau) = \beta_0, \text{ and} \tag{14}$$

$$\lim_{\tau\to\infty} f(\tau) \equiv \lim_{\tau\to\infty} i_c(\tau) = \beta_0 + \beta_1. \tag{15}$$

From the above limits, we observe that the instantaneous forward and spot rates can be approximated as the sum of the β_0 and β_1 , while β_0 is an approximation of the long-run rate (as known as, the steady-state level). Fitting the long-end of the term structure of the yield curve may be difficult as the convexity effects on bonds tend to pull down the yields on longer maturities (Gürkaynak et al., 2006). Gürkaynak et al. (2006) highlighted that the Nelson-Siegel specification tends to have forward rates asymptote too quickly to be able to capture the convexity effects at longer maturities.

The Nelson-Siegel model was later extended by Svensson (1995) through the inclusion of an additional exponential term which accounts for a second hump in the forward rate curve. The inclusion of this term increases the flexibility of the curve and improved the data fit. The functional form of the forward rate curve specified by Svensson (1995) is:

$$f(\tau) = \beta_0 + \beta_1 e^{-\lambda\tau} + \beta_2 \lambda\tau e^{-\lambda\tau} + \beta_3 \gamma\tau e^{-\gamma\tau}, \tag{16}$$

where $\theta = (\beta_0, \beta_1, \beta_2, \beta_3, \lambda, \gamma)$ is a vector of parameters to be estimated. Similarly, the location and size of the second hump is governed by β_3 and γ. Note that the Svensson model collapses to a Nelson-Siegel model if $\beta_3 = 0$. Integrating Equation 16 (with respect to τ) from 0 to τ, and dividing the result by τ, the outcome is the continuously compounded spot rate curve:

$$i_c(\tau) = \beta_0 + \beta_1\left(\frac{1-e^{-\lambda\tau}}{\lambda\tau}\right) + \beta_2\left(\frac{1-e^{-\lambda\tau}}{\lambda\tau} - e^{-\lambda\tau}\right) + \beta_3\left(\frac{1-e^{-\gamma\tau}}{\gamma\tau} - e^{-\gamma\tau}\right). \tag{17}$$

Similar to the Nelson-Siegel model, the Svensson model converges to similar limiting points at both ends of the curve. The estimation of the Svensson model relies on fitting data to Equation 16 to obtain the beta coefficients, λ and γ parameters.

4. DATA AND ESTIMATION ISSUES

4.1 Method of Estimation

In estimating the yield curve, the Svensson method was considered. The procedural method of estimation adopted in the study follows closely to that of Kladivko (2010).[1] The estimation of the parameters relies on the minimization of the weighted sum of squared deviations between the actual and predicted bond prices of coupon bonds:

$$\hat{\theta} = \arg\min_{\theta} \sum_{i=1}^{N}\left(\frac{P_i - \hat{P}_i}{P_i D_i^M}\right)^2, \tag{18}$$

where N is the number of observed bonds, P_i is the observed dirty price of the coupon bond, θ is the vector of parameters to be estimated, $\hat{P}_i$ is the estimated bond price which is obtain from the model spot rates, Equation 1 the discount function and Equation 4 the bond price formula. Similar to Kladivko (2010), the inverse of the product of observed bond prices and modified duration, $\left(1/P_i D_i^M\right)$, were adopted

[1] The Matlab codes developed by Kladivko (2010) were utilized for this paper.

as the optimization weight. The continuously compounded spot rates were obtained under the day count convention of *30/360,* for interest accrued.

The implementation of Equation 16 was conducted with *Lsqnonlin* in Matlab, a nonlinear least squares algorithm developed in Coleman and Li (1996). Due to its flexibility, *Lsqnonlin* allows for setting of the lower and upper bound of parameter(s) to be optimized, hence making it ideal for estimating parametric models of the yield curve. However, a drawback of the optimization algorithm *Lsqnonlin* is its sensitivity to the initial value of λ, as mentioned by Kladivko (2010). He advised that given the true value of λ, the algorithm converges robustly to the true values of β parameters of the parametric model of interest. From this, he concludes that the *Lsqnonlin* algorithm succeeds in finding the global minima. Despite these pros and cons, the initialization of the parameters of the models follows closely to that of Kladivko (2010) and Gürkaynak et al. (2006).

The estimation of parameters of the yield curve may suffer from abrupt changes in their values from one period to the next. Such changes were referred to as catastrophic jumps by Cairns and Pritchard (2001). In addressing catastrophic jumps in the estimated level component of the yield curve, β_0, Kladivko (2010) imposes a lower bound on the possible values that λ and γ may assume. Additionally, Kladivko (2010) restricted β_0 to be positive which is in line with the theory. These constraints give rise to restrictions on the parametric models as pointed out by Kladivko (2010). He further pointed out in his study that the restricted Nelson-Siegel model does not perform much different when compared to the unrestricted Nelson-Siegel model. However, unlike Kladivko (2010) who relies on daily data for his analysis, this study utilizes quarterly data on bond prices which makes it difficult to observe catastrophic jumps in the parameter estimates.

4.2 Data Set

The study utilizes quarterly market values of domestic GOJ bonds reported by domestic market participants for the period 2014Q1 to 2016Q4. This sample period was chosen because the data that were available prior to the selected period were perceived to be noisy in relation to the developments that took place in 2010 and 2012. During the first quarter of 2010, the GOJ conducted a restructuring of their debt portfolio. The restructuring of the government's portfolio was due primarily to the challenge in servicing the existing debts at the given maturities. As such there was a shift in most maturities to longer tenor. Similar actions were performed by the government in the first quarter of 2012. Since then, the government has reduced its participation in the domestic market significantly.

To date, the existing domestic bond market lags behind that of developed and transitional states as trades in these instruments are not captured in a formal trading system. In light of this, the market value reported by the domestic participants at the end of the quarters were used to extract the average bond prices. The data used in the study came from two primary sources: Financial Services Commission for information on nonbank financial institutions and Bank of Jamaica for information on deposit-taking institutions.

In improving the quality of the estimation, a data filtering process was developed. For the period under study, the following data cleaning was conducted:

1) Benchmark investment notes identified by the GOJ were utilized.[2]

2) Floating interest rate bonds were excluded since their use in estimating the yield curve is not straight-forward.

[2] Includes domestic issued JMD denominated securities that have a noncallable feature.

3) For each benchmark notes, bond prices that exceed two standard deviations about its mean were excluded from the analysis so as to minimize possible distortions in the data.

4) No adjustments for tax or coupon effects were made.

5) Bonds that were issued for more than one year and mature within six months are excluded as they distort the liquidity conditions in the market.

6) Bonds that were issued for less than six months that matures over one year were also excluded from the sample due to their liquidity conditions.

In total, 12 GOJ bonds' data were used for the period under study. In fitting the short end of the curve, the one month, three months and six months Treasury bill rates were utilized. The fitting of the short end reduces the likelihood of obtaining negative rates or extremely high rates which is important in the estimation process. A key advantage of the data reported is the richness of information collected.

5. ESTIMATION RESULTS

Using the above methodology, the Svensson yield curve was estimated for the period March 2014 to December 2016. The evolution of the estimated curve throughout the period was fairly stable as observed from the parameter estimates (see Figure 1).[3] The level parameter of the model fluctuated around a marginally improving trend within the bands of 8% and 19%. Except the third quarter 2014, the slope parameter of the model posited a slightly upward trend below the zero mark. Similarly,

[3] It was noted throughout the sample period that there were quarters in which the estimated results of Svensson model imply over parameterization (see Annex A). Alternatively, one may estimate a Nelson-Siegel model which was also considered by the study.

the curvature parameters (λ and γ) were slightly trending upward over the sample period. The interest rate spread between the 10-year and 1-year yields gently sloped upwards over the estimation horizon. At the long-end, the spread between 35-year and 10-year yields fluctuated around a relative downward sloping trend line. The interest rate spread between the 1-year and 10-year yields was highest for 2015Q3 where the corresponding spread at the long end of the curve was lowered.[4] This outturn to some extent reflected investors' preference along the maturity spectrum for the GOJ's domestic JMD issues. At the long end of the curve, interest rate spread was highest for 2015Q1 which corresponded to a decrease in the corresponding interest rate spread for the 1- to 10-years yields when compared to 2014Q4.[5] For the period 2014Q4, interest rate spreads for 1- to 10-years yields and 10- to 30- years yields recorded a positive quarterly growth, thus reflecting to some extent increased preference for higher yields across the entire maturity spectrum of the GOJ domestic JMD issue.[6] The flattening of the curve at the long end was most evident for 2014Q3 which reflected the minimum interest rate spread for 10- to 30-years yields over the sample period.

In sum, the estimated outputs throughout the sample period provided upward sloping yield curves.[7] The fit of the model to the observed sample data was most accurate as at end-2015 as displayed by the incorporated error measures.

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4 The 1- to 10-years spread on yields was 4.6% reflecting a 10.1% increased relative to 2015Q2 while the 10- to 30-years yields spread was 2.6% reflecting a 29.5% decline relative to the prior quarter.

5 The interest spread between the 10- to 30-years yields was 5.9% reflecting 12.7% increase while the 1- to 10-years interest rate spread was 4.1% reflecting a decline of 2.8 percent.

6 The interest spread between the 1- to 10-years and 10- to 30-years yields were 4.2% and 5.3% reflecting quarterly increases of 5.3% and 178.5%, respectively.

7 See Estrella and Trubin (2006).

Figure 1

ESTIMATED ZERO-COUPON YIELD CURVES

Continuously compounding

March 2014 to december 2016

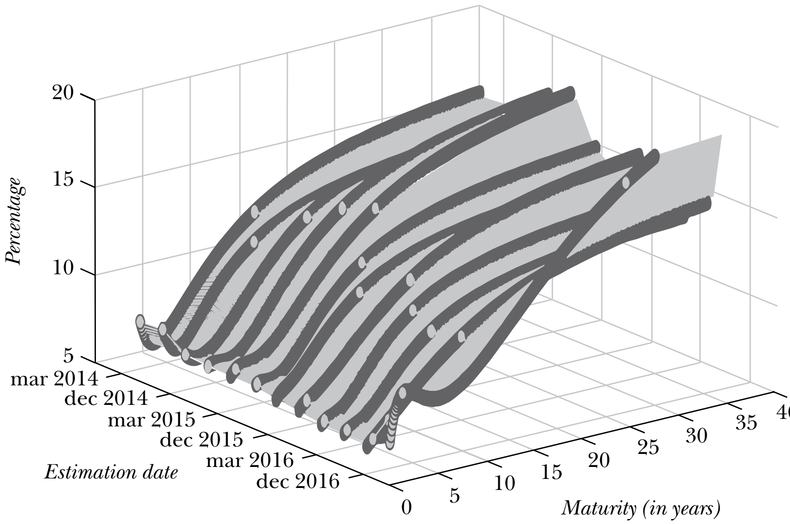
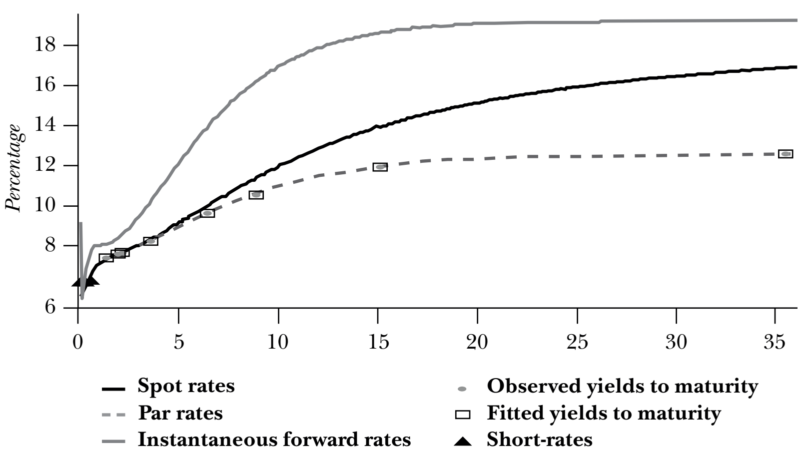


Figure 2

ESTIMATED OUTPUT FOR THE SVENSSON MODEL

As at the end-december 2015



As an example of the results, the estimated spot, instantaneous forward and par rates for December 2015 were captured by Figure 2. The rates are presented as annually compounded. There were eight government bonds available as at end-2015 with maturities ranging from approximately one year and four months to approximately thirty-five years.

As can be seen from Figure 2, the Svensson curve provides a fair fit of the term structure of the government's domestic debt. However, the fit of the curve was poorer at the short-end of the curve (less than one year) reflecting the idiosyncratic nature of these issues. For the one to five years maturity bucket, the fit of the 2019 8.5% coupon bond was the worst which appeared to be overpriced relative to the other bonds. The shape of the estimated spot rate curve was upward sloping for maturities over three years. At the short end, a U-shaped hump was evident. This suggests market participants' expectation of monetary easing by the central bank in the short term, (Bomfim, 2003).

Similar to Kladivko (2010), the mean absolute error (MAE), the root mean squared error (RSME) and the maximum absolute error (MaxAE) were used to assess the goodness of fit of the model:

19 $$\mathrm{RSME} = \sqrt{\frac{1}{n}\sum_{i=1}^{n}\left(y_i - \hat{y}_i\right)^2},$$

Table 1

ERROR MEASURES FOR ESTIMATED YIELD CURVE
Svensson estimated yield to maturity curve at the end of December 2015
Basis points

RSME	*MAE*	*MaxAE*
3.8	3.3	6.7

20 $$\mathrm{MaxAE} = \max_i \left\{\left|y_i - \hat{y}_i\right|\right\},\ \ i = 1, \ldots, n,$$

where n is the number of government bonds for a given settlement date, y_i is the observed yield to maturity, and $\hat{y}_i$ is the fitted yield to maturity. In calculating the error measures, the Treasury bill rates were excluded from the analysis.[8]

The estimated MaxAE which identifies the point of least best fit was associated with the 2018 7.75% coupon bond. The MaxAE for the estimated 2015Q4 zero-coupon curve reflected the overpricing of the 2018 7.75% coupon bond relative to the corresponding estimated output.

6. STRESS TESTING APPLICATION OF THE YIELD CURVE

The yield curve has many applications that are localized to the intended purposes. For example, inflation expectation which is of critical importance for monetary policy can be obtained from the yield curve. Additionally, Estrella and Trubin (2006) investigated the use of the yield curve as a forecasting tool in real time of macroeconomic conditions. The study employed a probabilistic model to capture the relation between key attributes of the curve (that is, the steepness of the curve) and the business cycle, for which they found that the yield curve was a good predictor of recessions.

Seminal work of Ho (1992) utilized nonparallel shifts in the yield curve as an approach for fixed income portfolio immunization. Ho (1992) investigated the impact of changes in selected rates along the curve on the pricing of fixed income securities. This approach is currently coined *key rate duration* (KRD) and is commonly used among financial market practitioners in developing hedging strategies for their portfolio holdings.

This paper applied the key rate model to GOJ's domestic sovereign portfolio within the context of assessing interest rate risk

[8] The exclusion of the error measures for Treasury bill rates was motivated by the poor fit of the curve at the short end. In addition, yields on Treasury bill were not collected in the sample.

exposure. Such applications involved shifting of the zero-coupon curve through selected key rates for the GOJ domestic JMD bond portfolio. With these key rates, one has the flexibility to conduct parallel and nonparallel shifts of the curve to provide richer analysis of bond price movements.

6.1 Key Rate Model

In this section, the KRD and the key rate convexity measures of interest rate risk are discussed along the lines of application for stress testing. The KRD as defined by Ho (1992) is a measure of the price sensitivity of a fixed income security to changes in selected spot rates along the yield curve. These rates are referred to as the key rates. Ho (1992), who pioneered the application of the KRD for fixed income portfolio, recommended 11 key rates: 1, 2, 3, 4, 5, 7, 9, 10, 15, 20 and 30 years to maturity. It is important to note that the choice of key rates along the yield curve is flexible in that one can choose any number of rates rate along the curve. The KRD measure is used by market practitioners to decompose portfolio returns, identify interest rate risk exposure, design active trading strategies and implement passive portfolio strategies such as portfolio immunization and index replication (Nawalkha et al., 2005).

The use of the key rate model is conditional on the assumption that any smooth change in the term structure of zero-coupon yields can be represented as a vector of changes in a number of properly chosen key rates. That is:

$$\Delta Y = \left[\Delta y(t_1), \Delta y(t_2), \ldots, \Delta y(t_m)\right], \tag{21}$$

where Y is the zero-coupon curve and $\Delta y(t_i)$ for $i = 1, 2, \ldots, m$ are the set of m key rates. Changes in all other interest rates are approximated by linear interpolation of the changes in the adjacent key rates. The shifting of a key rate along the zero-coupon curve, only impacts rates within the neighborhood of the selected key rate that are bounded to the right and the left by the closest key rates to our key rate of interest (Nawalkha et al., 2005). Rates outside of this bound will be unchanged. The shortest and longest key rates are bounded on one side, the shortest key rate is bounded to the right by the second key rate while the longest key rate is bounded to the left by the *m − 1st* key rate. Thus, shifting the shortest key rate by an amount x results in an equal amount in shifting rates to the left of the shortest key rate and a linear interpolation of the shift in rates to the right of the key rates that are bounded, while leaving rates above the bound unchanged. Similarly, shifting the longest key rate results in an equal shift of rates to the right of the longest key rate and linear interpolation of the shift in rates to the left of the longest key rate that are bounded, while leaving all other rates below the bound unchanged. A generic expression for the change in the interest rate for any given term t is written as:

22 $$\Delta y(t)=\begin{cases}\Delta y\left(t_{shortest}\right) & t \leq t_{shortest} \\ \Delta y\left(t_{longest}\right) & t \geq t_{longest} \\ \alpha \times \Delta y\left(t_{left}\right)+(1-\alpha)\times \Delta y\left(t_{right}\right) & else\end{cases},$$

where $y\left(t_{shortest}\right)$ and $y\left(t_{longest}\right)$ are the shortest and longest key rates, $y\left(t_{left}\right)$ and $y\left(t_{right}\right)$, with $t_{left} \leq t \leq t_{right}$, refers to the key rate adjacent (to the left and the right) to term t, and α and $(1-\alpha)$ are the coefficients of the linear interpolation, define as:

$$\alpha=\frac{t_{right}-t}{t_{right}-t_{left}},$$

$$1-\alpha=\frac{t-t_{left}}{t_{right}-t_{left}},$$

The set of key rate shifts can be used to evaluate the change in the price of fixed income securities. An infinitesimal shift in a given key rate, $\Delta y(t_i)$, results in an instantaneous price change given as:

23
$$\frac{\Delta P_i}{P} = -KRD_i \times \Delta y(t_i),$$

where KRD_i is the *i-th* KRD. So, the key rate is defined as the negative percentage change in the price of a given fixed income security resulting from the change in the *i-th* key rate:

24
$$KRD_i = -\frac{1}{P}\frac{\delta P}{\delta y(t_i)}.$$

Alternatively, the duration of the *i-th* key rate is defined as the negative of the elasticity of the price of a given fixed income security to the *i-th* key rate relative to the *i-th* key rate:

25
$$KRD_i = -\frac{e_{p,i}}{y(t_i)},$$

where $e_{p,i}$ is the elasticity of the price to the *i-th* key rate. The application of the key rate model is fairly straight forward. First, we calculate the KRD for each of our five key rates using the formula:

26
$$\frac{\delta P}{\delta y(t_i)} = \frac{\delta P}{\delta y(t)}\frac{\delta y(t)}{\delta y(t_i)} = \frac{CF_t \times t}{e^{y(t)\times t}}\frac{\delta y(t)}{\delta y(t_i)}.$$

By substituting Equation 22 into 18, we have:

27
$$KRD_i = t \times \frac{\delta y(t)}{\delta y(t_i)},$$

where t is the time to maturity. Observe that the KRD is an increasing function of time. Thus, key rates at the long end of the curve would have a greater responsiveness of price changes to interest rate changes.

The total price change resulting from all key rate changes is given as:

28
$$\begin{aligned}\Delta P &= \Delta P_1 + \Delta P_2 + \ldots + \Delta P_m \\ &= -\sum_{i=1}^{m} KRD_i \times \Delta y(t_i).\end{aligned}$$

The sum of the KRD measures from a simultaneous shift in all the key rates by the same amount results in the traditional duration of a given fixed income security. Thus, the KRD measure only account for the linear effect of key rate shifts. Under a non-infinitesinal shift in the term structure, the KRD framework is extended to account for second-order nonlinear effects of such shift. The nonlinear effect of the key rate shifts is called the key rate convexity (KRC) and is defined as:

29
$$KRC(i,j) = KRC(j,i) = \frac{1}{P}\frac{\delta^2 P}{\delta y(t_i)\delta y(t_j)}$$

for every pair *(i, j)*, of key rates. Similarly, the sum of the KRC measures from a simultaneous shift in all the key rates by the same amount results in the traditional convexity of a given fixed income security. The KRDs and KRCs of a portfolio can be obtained as the weighted average of the KRD and KRCs of the securities in the portfolio.

The following section discusses the selection of the key rates that will be used in our KRD model to conduct parallel and non-parallel shifts of the yield curve. Such shifts of the zero-coupon curve will be governed by scenario analyses that are acceptable industry practices.

6.2 Application of the Key Rate Model

The choice of key rates as pointed out by Zeballos (2011) is arbitrary owing mainly to the absence of unique fundamentals. In acknowledgment of this gap in the model framework, Nawalkha et al. (2005) proposed that the choice of key rates can be guided by the maturity structure of the portfolio under consideration. As such, the choice of key rates for this analysis will be guided by the structure of the government's domestic fixed income portfolio.

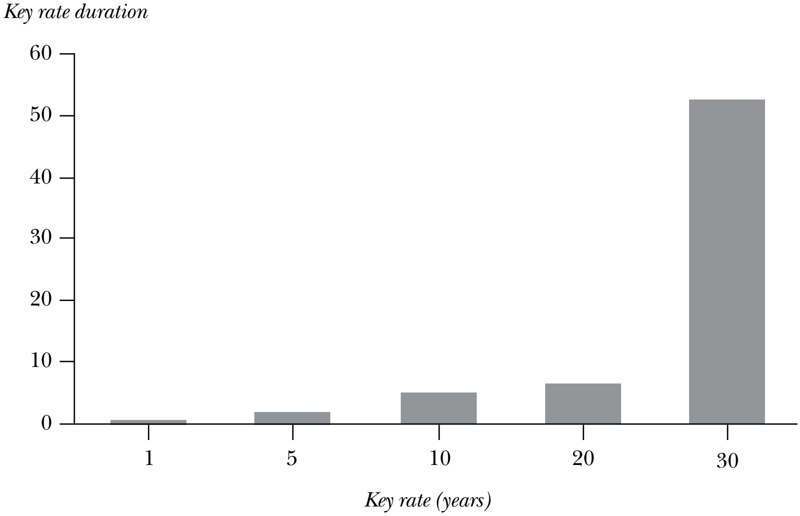
As at end March 2016, total outstanding JMD denominated government's issue was approximately 233 billion JMD in nominal value for fixed coupon bonds and 508 billion JMD in nominal value for variable coupon bonds which is unevenly distributed across 33 issues. This outstanding debt issue is sparsely distributed across the maturity spectrum of the yield curve. Approximately 50% of the outstanding debt matures within the next three years while 21% falls within the maturity range 20-35 years (see Figure A.2).

For this study five key rates were considered for varying reasons: the 1-year and 5-year were chosen as the major share of the government's domestic bond portfolio was at the short end; the 10-year key rate was reasonably viewed as a point along the curve ideal for conducting various shifts in the shape of the curve. For example, the butterfly shift of the curve, as well as a tilt of the curve, could be facilitated by fixing the 10-year key rate. The 20- and 30-year key rates provides useful analysis of the long end of the curve and are in line with the long-term maturity's share of the government's fixed income portfolio.

The result of our key rate application is presented in Figure 3. To calculate the KRD for the bond portfolio a shift of 100 basis points was applied to each of the key rates. Then, for each key a weight was assigned to each maturity conditional on the portfolio maturity spectrum. So, for example, rates that had time to maturity of one year or less were assigned a weight that represents the share of nominal issues that mature within one year. Likewise, rates one to two years was assigned a weight of nominal issues that mature one to two years.

Figure 3

KEY RATE DURATION



As evident in Figure 3, the portfolio has larger expositions over the medium to long-term. Specifically, the exposition for the 30-year key rate dominates the bond portfolio followed by the 20-year key rate.[9] This means that the bond portfolio is more sensitive to changes in the long end of the yield curve. Zeballos (2011) pointed out in a recent study that a concentration in the KRD at the long end of the term structure may indicate an expectation of the flattening of the yield curve.[10]

9 A KRD of 50 for the 30-year key rate means that a 100 basis points change in the 30-year key rate would lead to 50 percent reduction in the weighted aggregated value of the GOJ domestic JMD portfolio cash flows that have a maturity period greater than 20 years.

10 Similarly, the KRC for the bond portfolio was also calculated. The result of the KRC was in some sense similar to the outcome

6.3 Stress Testing Application of Yield Curve Shifts

As part of the Bank's interest rate stress test, scenario shifts in the yield curve are considered. This paper utilizes key rates to conduct parallel and nonparallel shifts in the yield curve. For a parallel shift in the yield curve, equal shifts in the selected key rates are considered. Nonparallel shifts in the yield curve amount to unequal shifts in the key rates. Specifically, an upward tilt of the yield curve at the 10-year key rate is achievable through an upward shift in key rates to the left of the 10-year key rate while simultaneously shifting the key rates to the right downwards. In the case of the domestic fixed income sovereign issues, four cases are considered for illustration: *1)* a parallel upward shift of the yield curve; *2)* a flattening of the curve at the short end up to 10-year; *3)* an increase in premiums for medium tenors; and *4)* a steepening of the curve at the long end of the maturity spectrum. The assessment of each scenario will be conducted based on changes of stress levels of 20%, 50% and 100% in the yields, respectively.

6.3.1 An Upward Parallel Shift of the Yield Curve

A parallel shift of the curve is supported by the notion of investors requiring equal premiums across the term structure due to higher perceived risk of government's ability to repay its debt. Such shift of the curve is accomplished by increasing the key rates by similar amount. The study considered 20%, 50% and 100% increases in the key rates simultaneously across the estimated term structure. The new yield curve was then used to evaluate fair value losses[11] for portfolio holding of deposit-taking institutions (DTIs), securities dealers and insurance companies.[12] The

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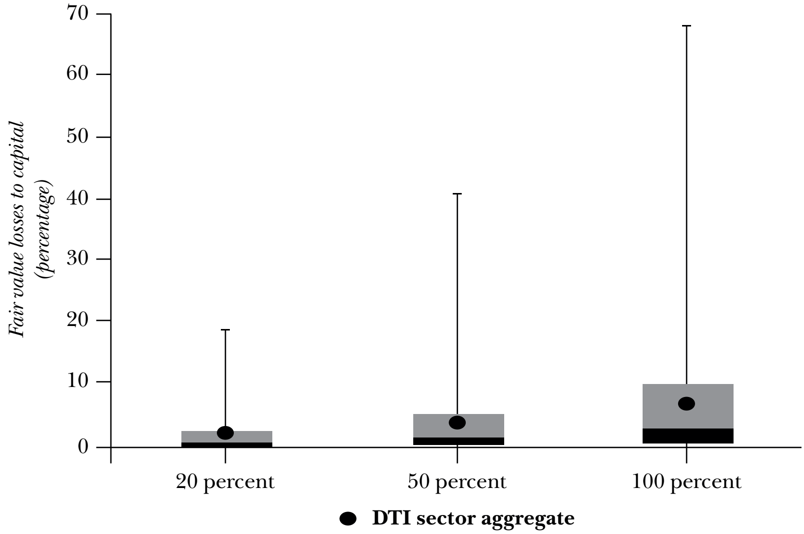
of the portfolio's KRD and are not included in the analysis for ease of explanation.

[11] Fair value loss is defined as the difference in value of GOJ domestic JMD portfolio holdings resulting from changes in yields.

[12] Currently, the deposit-taking subsector comprises of six commercial banks, three building societies and two merchant banks.

Figure 4

BOX-PLOT OF THE RATIO OF FAIR VALUE LOSSES TO CAPITAL FOR THE DEPOSIT-TAKING INSTITUTIONS SECTOR FOR PARALLEL SHIFTS OF THE YIELD CURVE



results of the parallel shift of the curve showed an impairment to the capital base of DTIs of 16.2% resulting from a 100 shock to the yield curve (see Figure 4).[13]

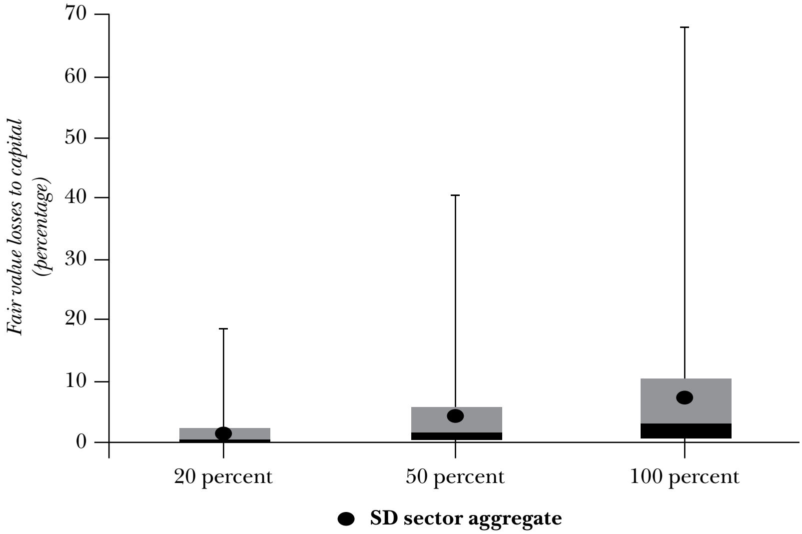
A 20% increase in the term structure had a marginal impact on the fair value losses of the DTI sector (3.8% loss in capital) while at a 50% shock levels, impairments to capital were 9% (see Table B.1 in Annex B). The impact of the 100% shock threshold level on individual institutions within the DTI sector resulted in no significant impairment to their capital adequacy ratio; hence, indicating that the DTI sector is adequately

These institutions account for approximately 50% of the total financial system's assets.

[13] Impairment to capital for each subsector is defined as the fair value loss divided by total accounting capital holding.

Figure 5

BOX-PLOT OF THE RATIO OF FAIR VALUE LOSSES TO CAPITAL FOR THE SECURITIES DEALER SECTOR FOR PARALLEL SHIFTS OF THE YIELD CURVE



capitalized to withstand such shocks in the yields on government's domestic issues.

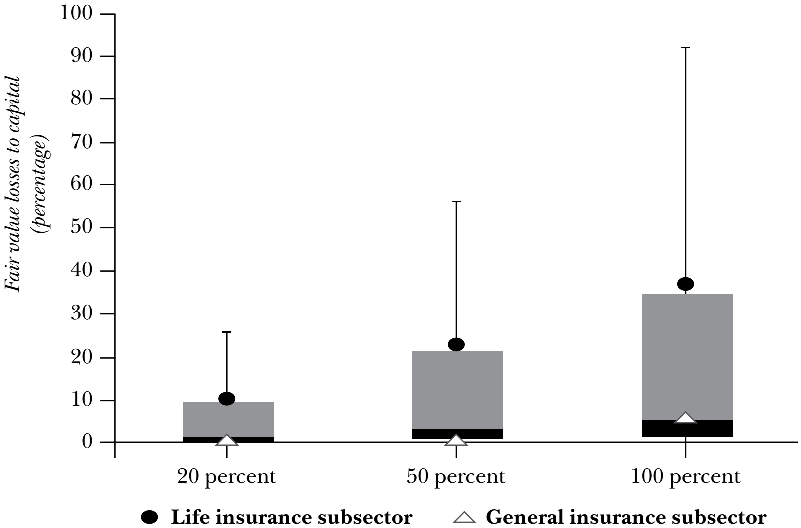
The result of the analysis revealed that securities dealers were less susceptible to parallel shifts of the curve than DTIs. The sector's impairment to capital from a 100% upward shift of the term structure was 7.5% (see Table B.1 in Annex B). A 20% increase in the term structure would result in an impairment to securities dealers' capital of 1.9% (see Figure 5), while a 50% increase in rates resulted in impairment of 4.3 percent.

At the 50% shock level, one institution fell below the capital adequacy ratio prudential minimum level of 10%. The outcome was unchanged at the 100-shock level where one institution fell below the capital adequacy ratio prudential minimum level.

An assessment of the insurance industry revealed that fair value losses from a 100 increase in rates across the term

Figure 6

BOX-PLOT OF THE RATIO OF FAIR VALUE LOSSES TO CAPITAL FOR THE INSURANCE SECTOR FOR PARALLEL SHIFTS OF THE YIELD CURVE



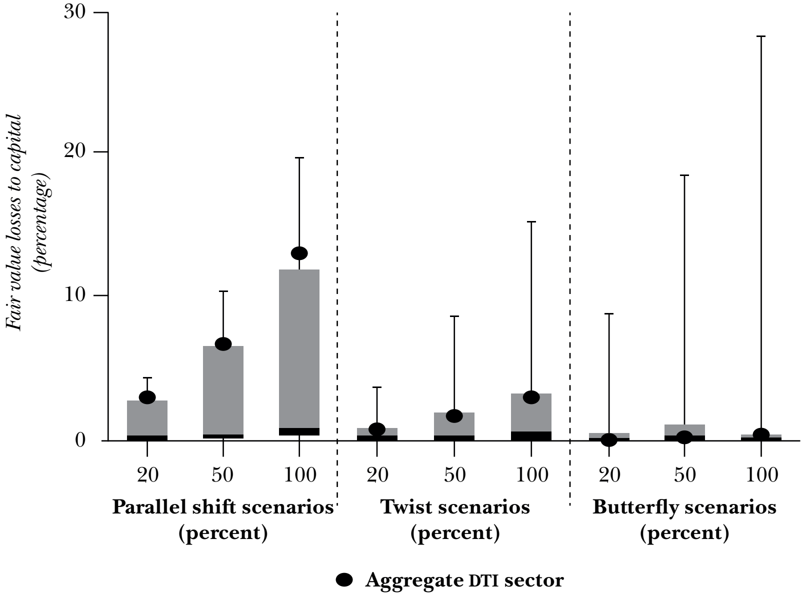
structure accounted for 37.4% of the life insurance subsector capital base (Figure 6). Exposure to the general insurance subsector, on the other hand, was less than 10% of its capital base (see Table B.1 in Annex B). At the 100% shock level, fair value losses across all three sectors of the market was highest for the insurance sector (specifically the life insurance subsector which accounted for 41.6% of total losses of 49.4 billion JMD).

6.3.2 Flattening of the Yield Curve at the Short End

A hypothetical flattening of the yield curve was considered, in which the 1-year key rate increased by 20%, 50% and 100%, respectively. Such movement in the curve would result in greater impact on portfolios holdings that are concentrated within maturities of up to five years. The outcome of the assessment showed that DTIs were more susceptible to the flattening of

Figure 7

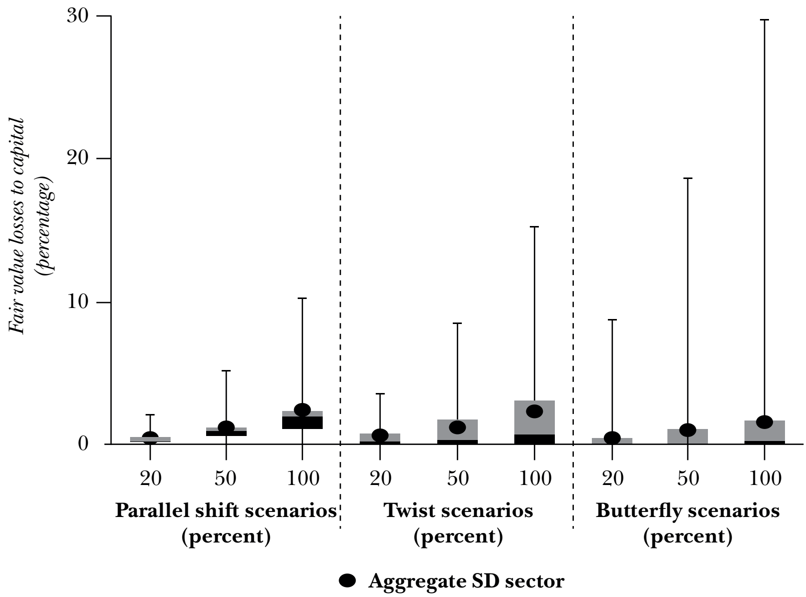
BOX-PLOT OF THE RATIO OF FAIR VALUE LOSSES TO CAPITAL FOR THE DEPOSIT-TAKING INSTITUTION SECTOR FOR NONPARALLEL SHIFTS OF THE YIELD CURVE



the curve at the short end than securities dealers. At the 100% shock level fair value losses for DTIs amounted to 12.7% of their capital base while securities dealers amounted 2.6% of their capital base (see Figures 7 and 8 and Table B.1). Similarly, life insurance subsector was more exposed to the flattening of the curve at the short end when compared to the general insurance subsector for the insurance sector (see Figure 9). At the 100% shock level fair value losses for life and general insurance subsectors were 3.1% and 2.2% of their capital base, respectively. In addition, across the market, the DTI sector had the greatest exposure to the stress testing of the short end of the curve followed by the life insurance subsector. Evidently, the outcome

Figure 8

BOX-PLOT OF THE RATIO OF FAIR VALUE LOSSES TO CAPITAL FOR THE SECURITIES DEALERS SECTOR FOR NONPARALLEL SHIFTS OF THE YIELD CURVE



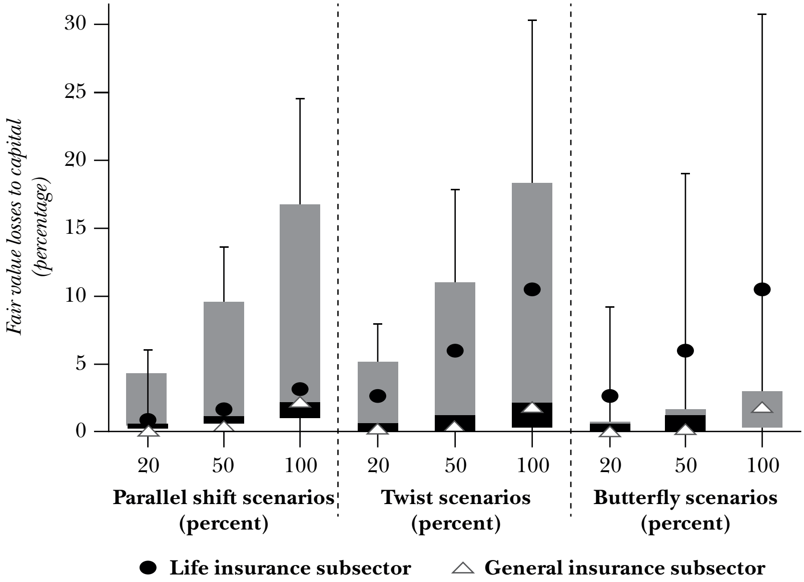
of the flattening of the curve was lower than that of a parallel shift of the curve.

6.3.3 An Increase in Premiums for Medium Tenures along the Curve

A hypothetical increase in yields along the medium-term tenures (for example, five years to ten years) of the curve was considered as an increase in the demand for premiums along these tenors by investors. To simulate such changes in the yield curve the 10-year key rate was adjusted upwards at the respective shock levels. The adjustment in the 5-year key rate would impact yields that are greater than the 5-year key rate up to the 10-year key rate and above the 10-year key rate but less than the 20-year key rate.

Figure 9

BOX-PLOT OF THE RATIO OF FAIR VALUE LOSSES TO CAPITAL
FOR THE INSURANCE SECTOR
FOR NONPARALLEL SHIFTS OF THE YIELD CURVE



The relative fair value exposure to capital for such movement along the curve was largest for the insurance sector across the market. At the 100% shock level, fair value losses from such movement along the curve was 10.4% of capital for the life insurance subsector and 2% for the general insurance subsector (see Figure 9, and Table B.1).

While for the DTIs and securities dealers, such movement along the curve would result in lower exposure when compared to a flattening of the curve at the short end. At the 100% shock level, fair value losses relative to capital were 2.8% and 2.4% for DTIs and securities dealers, respectively (see Figures 7 and 8).

6.3.4 A Steepening of the Curve at the Long End of the Maturity Spectrum

A hypothetical increase in yields along the long end (that is, above 10 years) of the curve was considered reflecting an increase in uncertainty of long-term macroeconomic conditions by investors. To simulate such movements in the yield curve, the 20-year and 30-year key rates were stressed at the respective shock levels. Relative to prior segmented shifts along the curve, exposures for the life insurance subsector was largest for shifts at the long end of the yield curve. At the 100% shock level, fair value losses from such movement along the curve was 10.4% of capital for the life insurance subsector (see Figure 8 and 9, and Table B.1). Conversely, relative to prior segmented shifts along the yield curve, exposures for DTIs and securities dealers were smallest for shift at the long end of the maturity spectrum. At the 100% stress level, fair value losses relative to capital were 0.3% for DTIs and 1.7% for the securities dealers sector (see Figures 7 and 9).

From the respected shifts of the yield curve, it was observed that a parallel shift of the curve would have the largest impact on the fair value of the portfolio holdings of GOJ domestic securities across the respective sectors in the above analysis. In relation to nonparallel shifts of the yield curve, the results of the analysis were to some extent consistent with the fundamental market practice of the respective sectoral market participants. The life insurance subsector was more vulnerable to the medium to the long end of the maturity spectrum which is reflective of the appetite of their investment horizon. The DTIs, securities dealers and the general insurance subsector, on the other hand, were more vulnerable to the short to medium term segment of the yield curve.

7. CONCLUSION

This paper estimated the GOJ domestic yield curves from 2014 to 2016 at a quarterly frequency. The estimation of the curves was based on the Svensson model. The model fits the GOJ bond price data well without being overparameterized and, thus, provides a consistent picture of GOJ's domestic yield curve evolution. The results from the estimation of the GOJ zero-coupon spot rate curve show upward sloping yield curve. With the exception of 2014Q4, investors' preferences along the curve vary inversely across the 1- to 10-years and 10- to 30-years segments maturity spectrum of the GOJ domestic JMD debt portfolio.

Additionally, the estimated yield curve was utilized in an interest rate risk analysis for selected financial market participant sectors in Jamaica. As a risk assessment exercise, the study investigated the impact of parallel and nonparallel shifts of the yield curve on the portfolio holdings of selected domestic financial market participant sectors. The approach of the study relies on the KRD model for interest rate risk management. The choice of the KRD model was motivated by nonparallel shift scenarios for the yield curve.

The results from a parallel shift of the estimated yield curve showed that the life insurance subsector was more exposed to such movements in GOJ domestic bond yields relative to other market participant groups. In relation to nonparallel shifts of the curve, DTIs, securities dealers and general insurance subsector were more vulnerable to shifts in short to medium terms segment of the yield curve. The life insurance subsector was more exposed to the medium to the long end of the yield curve. The results of the assessment provide useful insights on the financial market structure, which was consistent with market expectation on the investment horizon for these participants.

The key rate model is a very useful tool for hedging interest rate risk and is used by market participants along with other tools. In light of the model's application, there are limitations to its use. Firstly, the choice of key rates is somewhat subjective. Thus, the model offers no guidance on the choice of the risk factor to be used despite its importance. As a circumvention to this shortcoming of the model, different numbers and choices of key rates may be selected conditional on the maturity structure of the portfolio under consideration.

Secondly, the shift in the individual key rates provides an implausible yield curve shape. Further, the shift in the key rates assumes strong correlation of the neighboring rates which may not always be the case. In addressing this shortcoming of the model, Johnson and Meyer (1989) proposed the partial derivative approach. This approach assumes that the forward rate curve is split up into many linear segments and all forward rates within each segment are assumed to change in a parallel way. Under it each forward rate affects the present value of all the cash flows occurring within or after the term of the forward rate.

Lastly, the key rate model does not take into account past movements in past yield curves hence making the model inefficient in describing the dynamics of term structure because historical volatilities of interest rates provide useful information.

ANNEXES

Annex A

Figure A.1

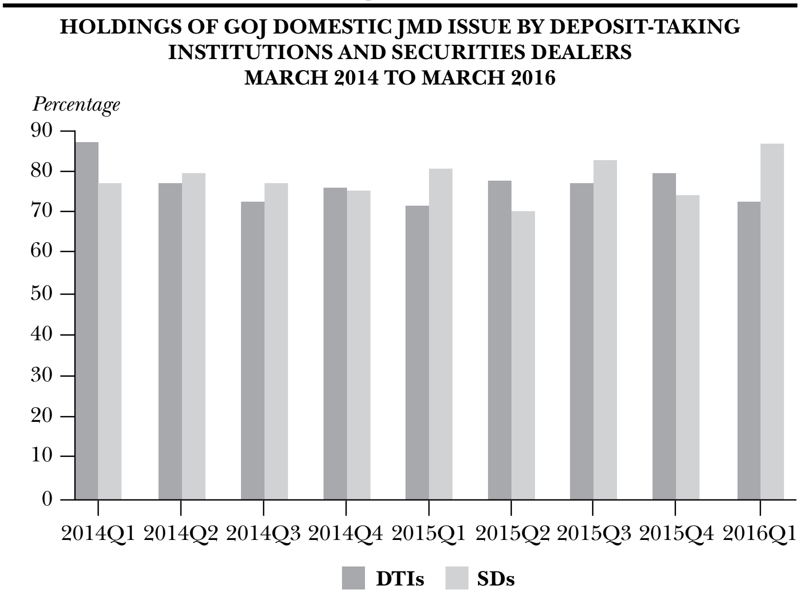


Figure A.2

DISAGGREGATION OF THE SHARE OF GOJ DOMESTIC JMD ISSUE BY MATURITY AS AT DECEMBER 2016

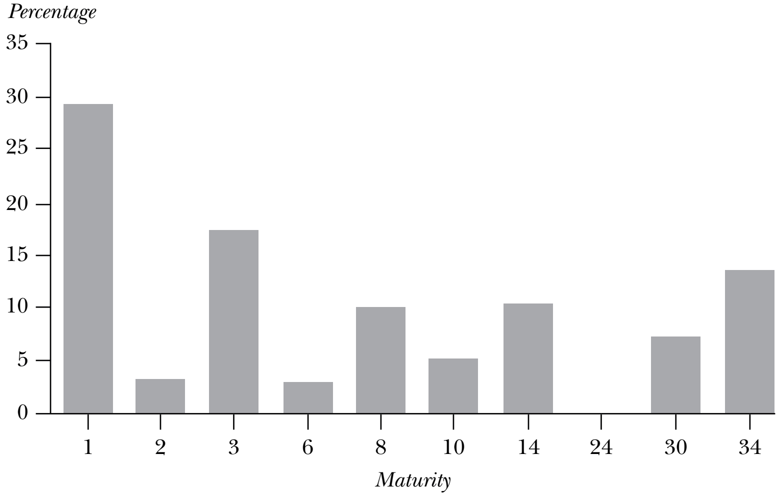


Table A.1

PARAMETER OUTPUT

Estimated Parameters for the Period 2014Q1-2016Q4

Actual values

Date	β_0	β_1	β_2	β_3	λ	γ
2014Q1	0.17	−0.13	−0.15	0.16	0.41	3.81
2014Q2	0.15	−0.11	−0.19	0.16	0.64	2.40
2014Q3	0.08	−0.00	−18.86	18.96	0.18	0.18
2014Q4	0.19	−0.13	28.22	−28.40	0.59	0.59
2015Q1	0.20	−0.12	−0.12	−0.33	2.52	0.38
2015Q2	0.17	−0.11	−22.06	21.93	0.74	0.74
2015Q3	0.15	−0.08	−0.09	−0.19	3.71	0.60
2015Q4	0.17	−0.12	−0.07	−0.05	0.23	2.63
2016Q1	0.14	−0.10	−16.9	−16.81	0.94	0.95
2016Q2	0.15	−0.10	−9.18	9.08	0.79	0.80
2016Q3	0.17	−0.10	−0.09	−0.02	0.28	6.40
2016Q4	0.27	−0.15	−0.36	−0.51	2.92	0.23

Annex B

Table B.1

FAIR VALUE LOSSES RELATIVE TO CAPITAL FROM KEY RATE SHIFTS OF THE ESTIMATED YIELD CURVE AS AT DECEMBER 2016

		Shock levels (%)		
		20	*50*	*100*
Parallel upward shift of the curve	DTIs	3.8	9.0	16.2
	SDs	1.9	4.3	7.5
	LIs	11.1	23.4	37.4
	GIs	0.5	1.2	6.2
Flattening of the curve at the shortend	DTIs	2.9	6.9	12.7
	SDs	0.5	1.2	2.6
	LIs	0.7	1.6	3.1
	GIs	0.2	0.4	2.2
Increase in medium term tenures along the curve	DTIs	0.7	1.6	2.8
	SDs	0.6	1.3	2.4
	LIs	2.6	5.9	10.4
	GIs	0.2	0.4	2.0
Steepening of the curve at the long-end	DTIs	0.1	0.2	0.3
	SDs	0.5	1.1	1.7
	LIs	2.6	5.9	10.4
	GIs	0.2	0.4	2.0

Note: DTIs stands for deposit-taking institutions sector; SDs for securities dealers sector; LIs for life insurance subsector; and GIs for general insurance subsector.

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Inflation and Public Debt

José Pablo Barquero Romero
Kerry Loaiza Marín

Abstract

This article aims to determine if a deterioration in public finances, understood as an increase in public debt, tends to increase inflation. We study the relation between public debt, economic growth, money supply growth and inflation. To do this we follow the methodology proposed by Kwon et al. (2009), who perform a panel data estimation using a sample of net debtor countries. We find that for countries whose public debt is already high, further increases in public debt are inflationary.

Keywords: fiscal policy, monetary policy.
JEL classification: E60, E63.

1. INTRODUCTION

Inflation is considered a monetary phenomenon, meaning its control is conditional on monetary policy. The quantity theory of money argues that inflation is solely determined by changes in the relative supply of money and goods. Thus, policies aimed at reducing inflation have focused on constraining monetary expansion to keep it in line with the expansion in nominal income. Nevertheless, it has been propounded

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that money demand also depends on inflation expectations, suggesting that a purely monetary effort at reducing inflation might not the only factor worth considering. As a consequence, growing attention has been given to the role of fiscal policy in determining inflation.

The seminal work of Sargent and Wallace (1981) states that the effectiveness of monetary policy in controlling inflation depends critically on its coordination with fiscal policy. The authors argue that, even when the traditional connection between money and the price level holds, tight monetary policy could lead to increases in inflation. This is due to the fact that, with the demand for government bonds given and in the absence of changes in future fiscal policy, a part of government obligations would have to be covered by seignorage at some point in the future.

A similar school of thought lies behind the so-called fiscal theory of the price level (FTPL). This not only focuses on seignorage financing, but also on traditional analysis of the fiscal impact, particularly on Keynesian aggregate demand type factors, public wage spillovers to private sector wages, and taxes affecting marginal costs and private consumption (Elmendor and Mankiw, 1999).

The FTLP also identifies the wealth effect of government debt as an additional channel of fiscal influence on inflation. This theory contends that increased government debt adds to household wealth and, therefore, to demand for goods and services, leading to price pressures (Buiter, 1999; Niepelt, 2004; Sims, 1994; Woodford, 1994 and 2001; Loyo, 1999; Christiano and Fitzgerald, 2000; Canzoneri et al., 2001; Cochrane, 2001 and 2005; Gordon and Leeper, 2002). The higher size of the debt also results in higher sovereign risk premiums being charged by government creditors, which can increase interest rates in the economy as a whole and unleash the well-known crowding out effect with its accompanying impact on macroeconomic stability.

High levels of public debt and recurring fiscal deficits in Costa Rica might generate inflationary pressures under the reasoning mentioned above. Thus, in order to foster domestic stability, it is necessary to understand the link between public finances and inflation. Furthermore, the possible implementation of fiscal reforms to redress Costa Rica's public finances could affect the aforementioned relation due to its impact on economic growth as well as on the fiscal deficit and the size of government debt. A lack of information on fiscal reforms in Costa Rica limits the analysis, making it necessary to use approaches based on other fiscal variables such as government debt.

This paper aims to determine if a deterioration in public finances, understood as an increase in public debt, tends to increase inflation. We study the relation between public debt, economic growth, money supply growth and inflation. To do this we follow the methodology proposed by Kwon *et al.* (2009), who perform a panel data estimation using a sample of net debtor's countries. We find that for countries whose public debt is already high, further increases in public debt are inflationary.

Section 2 describes the literature on the relation between public finances and inflation. Section 3 then explains the methodology used, the theoretical and econometric approaches, the data employed, and the estimation process followed. The outcomes are presented in Section 4. Finally, Section 5 gives the main conclusions.

2. ANTECEDENTS

The size and persistence of fiscal deficits together with their variations over time and across countries is a topic that has drawn attention in theoretical and empirical fields, above all, with respect to the causes of these persistent deficits and their corresponding impact on public debt. Such deficits are considered a cause of money supply growth, persistent inflation and macroeconomic instability (Saleh and Harvie, 2005; Catão and Terrones, 2005; Tekin-Koru and Özemen, 2003; Hossain and Chowdhury, 1998). Tanzi (1993) even argued that, especially in developing countries, the public sector, far from being a balancing factor, has contributed to generating larger macroeconomic imbalances. Along the same lines, Fisher and Easterly (1990) point to the fact that rapid inflation is almost always a fiscal phenomenon and that controlling inflation requires monetary and fiscal policy coordination.

Ghura and Hadjimichael (1996) demonstrate that there is an inverse relation between economic growth and macroeconomic stability measured by the inflation rate and the fiscal deficit as a proportion of gross domestic product (GDP). Empirical evidence suggests persistent deficits are without any ambiguity whatsoever detrimental to economic growth (Easterly et al., 1994). Nevertheless, other studies find that only inflation in excess of 10% to 20% poses any real threats to economic growth (Gylfason and Herbertsson, 2001; Loungani and Swagel, 2003). Even so, there is no doubt that price stability–that is, low and stable inflation–is a basic requirement for sustained economic growth, while fiscal deficits and public debt should be maintained at levels in line with other macroeconomic targets, including controlling inflation (Easterly et al., 1994).

Despite the large body of research on the relation between debt, money, and inflation, no theoretical or empirical consensus exists on the exact economic consequences of large budget deficits on inflation (Darrat, 2000; Narayan et al., 2006). According to Sargent and Wallace (1981), inflation is associated to the way budget deficits are financed, it means, the extent to which deficits are monetized. The degree to which monetary policy is independent and budget policy dependent o vice versa is key to knowing whether fiscal deficits would lead to higher rates of inflation (Sargent and Wallace, 1981).

Elaborating on this theme, Vamvoukas (1998) and Saleh and Harvie (2005) mention the existence of two transmission channels of the deficit to inflation. First, when a central bank purchases government bonds, which increases high-powered money, the money supply, and thereby the price level. Second, when deficits put an upward pressure on interest rates that then require an increase in the money supply to keep them stable, in which case deficits cause inflation by encouraging higher rates of monetary growth. As Vamvoukas (1998) posits, in a world without a Ricardian regime,[1] increases in the real value of bond assets increase perceived private wealth that, added to income obtained from interest rates, makes bond holders feel richer, inducing them to increase their consumption spending. This leads to higher national income, but, this expansion of national income leads to an increase in the demand for money and with that inflation (Keynesian perspective).

In contrast, Barro (1996) and other proponents of the Ricardian equivalence contend that government deficits do not matter given that current tax cuts will be financed by proportionate future tax hikes, ensuring that the government deficit does not affect the economy. As opposed to the Keynesian viewpoint, current tax cuts and future tax hikes will offset each other, meaning tax cuts will not make economic agents wealthier and do not encourage them to increase their consumption of goods and services. Hence, fiscal deficits do not matter because they do have any effect on aggregate demand, interest rates, and the price level. For Barro (1996) the net value of private sector wealth remains unchanged by taxes or debt financing, which is the reason why deficits do not cause inflation. On the contrary, deficits would be the result of inflation.

Another channel by which a government deficit might directly affect inflation is through the output gap. The reasoning behind this is that the public sector also demands goods and services produced by the private sector. Nevertheless, such

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[1] In a non-Ricardian regime agents do not believe that changes in the shape or size of government financing lead to corresponding future alterations. Agents do not, therefore, include government budget restrictions in their decision-making. This means the method used to finance government expenditure affects intertemporal consumption decisions and, therefore, aggregate demand.

effect can be positive or negative depending on the type of public expenditure. For instance, if the public deficit is the result of greater current expenditure on goods and services, the expected effect would be positive. However, if said expenditure is used to construct infrastructure, the effect could be negative (at least over the long run), given that it would tend to improve productivity and lower production costs for the private sector.

In a similar way to theory, empirical evidence does not exhibit consensus with respect to the direction of the causal relation between inflation, fiscal deficit, and money. Choudhary and Parai (1991) find that budget deficits as well as the rate of growth of money supply have a significant impact on inflation in Peru. Meanwhile, Hondroyiannis and Papapetrou (1997) find bidirectional causality between inflation and budget deficits in Greece. In the case of Turkey, Metin (2012) finds that fiscal expansion is a determining factor for inflation and that budget deficits (as well as real income growth and debt monetization) significantly affect inflation. Likewise, for the case of South Africa, Anoruo (2003) shows evidence that deficits have a positive impact on the growth rate of money supply and inflation.

Catão and Terrones (2005) find a strong positive association between deficits and inflation among high-inflation and developing country groups. On the other hand, for low-inflation advanced economies the authors do not find a relation between budget deficits and inflation. Wolde-Rufael (2008) obtains empirical evidence for a long-run cointegrating relation between inflation, money, and budget deficits in Ethiopia, with a unidirectional Granger causality running from money supply to inflation and budget deficits to inflation, while monetary policy does not seem to have any impact on the growth of money supply. Meanwhile, Barro (1989), Abizadeh et al. (1996), Vieira (2000), and Wray (2005) argue that the inflation-deficit nexus does not exist because larger deficits do not cause inflation.

Moving away from a budget deficit focused approach, Castro et al. (2003) estimate the degree of interdependence between fiscal and monetary policies in developed countries by using government debt in itself rather than the budget deficit. These authors find that debt plays a minor role in determining the price level in developed countries. Along the same lines, Kwon et al. (2009) use panel dataset, separating developed and developing countries, as well as net debtor or net credit countries based on their balance of payments data and classification of the World Economic Outlook 2005 (IMF, 2005). They find that the relation between debt and inflation is statistically significant and strong in indebted developing countries, weak in other developing countries and generally not valid in developed economies (Kwon et al., 2009). The outcomes of Castro et al. (2003), as well as those of Kwon et al. (2009), are in line with the fiscal theory of the price level (FTPL) described previously.

3. RESEARCH METHODOLOGY

This paper follows the methodology of Kwon et al. (2009) and uses a panel dataset of annual data for 52 countries spanning 1965 to 2014. We employ a forward-looking model of inflation that is based on rational expectations, Cagan-type money demand[2] and a non-Ricardian[3] regime that takes government bonds as net wealth.

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[2] Cagan-type money demand takes the following form: $m_t^d - p_t = -\alpha E_t\left(p_{t+1} - p_t\right)$, where m_t^d is the log of nominal money held at the end of period t, p is the log of the price level, and α is the semielasticity of real money demand with respect to expected inflation. The exclusion of real variables such as output and interest rates is justified by arguing that during hyperinflation expected inflation cancels out all other influences on the demand for money (Cagan, 1956).

[3] As mentioned previously, in a non-Ricardian regime agents do not take into consideration government budget constraints because from their viewpoint current tax cuts or hikes will not necessarily be offset by any equivalent future taxes imposed by the government. Thus, the method used to finance government expenditure affects wealth and therefore agents' intertemporal consumption decisions and aggregate demand.

A functional relation can be derived for the price level with respect to debt, money, and real GDP, which is written in the following form (see Annex 4 for its foundations):

$$P_t = \left(\frac{M_t + \delta B_t}{\gamma(i) w} \right). \tag{1}$$

where,

$$\gamma(i) = \beta \left(\frac{1 + i_t}{i_t} \right) + \alpha\delta. \tag{1a}$$

and, P is the price; M, money; B, government debt; w, real income or wealth; α and β are functions of the structural parameters of the household maximization problem; i, yields on the debt; and δ is a part of the government debt that is not guaranteed by the government's current or future primary surpluses.

Equation 1 nests the quantitative theory of money and the unpleasant monetary arithmetic[4] of Sargent and Wallace (1981). The price level is proportionate to the monetary aggregate broadly defined as $M_t + \delta B_t$, which is the sum of high-powered money demanded by agents for transactions and by the government for debt monetization, with δ reflecting the extent of the budget deficit, that is, the coordination between monetary and fiscal policy.

To clarify Equation 1, suppose the government pursues a policy of not monetizing its debt and runs a balanced budget over the long term. The monetization factor δ then reduces to zero and the equation simplifies into the conventional quantity theory of money. Along the same lines, if fiscal policy is undertaken flexibly in ways to keep the debt-to-GDP ratio fixed all the time, then the monetization factor will remain at

[4] The purpose of that paper was to argue that even when monetarist assumptions are satisfied, the list of items monetary policy cannot control should be widened to include inflation.

zero and public debt will have no impact on the price level. Alternatively, if the policy arrangement is full monetization of public debt, then δ becomes 1, meaning that the issuance of the public debt influences inflation as strongly as money supply does. In reality, this parameter should vary between 0 and 1, with the exact scale depending on the capacity and willingness of the government to service public debt, which, in turn, depends on the debt size, policy credibility, and institutional and political constraints.

Although, following Kwon et al. (2009), the wealth effect of government debt is not explicitly included, as set forth by the FTPL (Leeper and Yun, 2006), Equation 1 is still consistent with the predictions of the FTPL. However, this means that the establishment of a positive significant relation between public debt and the price level does not necessarily answer whether it stems from debt monetization as suggested by Sargent and Wallace (1981) or the wealth effects postulated by the FTPL.

For Equation 1 the following generalized prices function can be used:

$$P_t = f\left(X_t\right) = f\left(M_t, B_t, w_t\right), \text{ where } f_1 > 0, f_2 > 0, f_3 < 0. \quad (2)$$

Equation 2 can be log-linearized around the equilibrium values X^* to obtain the following specification:

$$\log P_t = f\left(X_t^*\right) + X_t^* f'\left(X_t^*\right) x_t, \text{ where } x_t = \log X_t - \log X_t^*.$$

Therefore,

$$p_t = f\left(X_t^*\right) - \log P_t^* + X_t^* f'\left(X_t^*\right) x_t, \text{ where } p_t = \log P_t - \log P_t^*. \quad (3)$$

3.1 Empirical Approach

The previous transformation establishes a linear relation between inflation and increases in money supply, public debt, and output. Equation 3 can be modified to a dynamic setting that includes a process of restoration to the equilibrium (Hendry et al., 1984):

$$\hat{p}_t = \alpha \hat{p}_{t-1} + \beta_1 \hat{m}_t + \beta_2 \hat{b}_t - \beta_3 \hat{w}_t, \tag{4}$$

where $\hat{p}, \hat{m}, \hat{b}$ and $\hat{w}$ represent deviations from equilibrium values in logarithms of prices, money, debt, and real income, respectively.

To model Equation 4 we used a panel dataset that allows for variability of individual countries while preserving the dynamics of adjustment within countries. The following dynamic model, the formulation of which is based on Equation 4, is employed:

$$\begin{aligned} d\log cpi_{it} &= \alpha d\log cpi_{it-1} + \beta_1 d\log money_{it} + \\ &+ \beta_2 d\log debt_{it} - \beta_3 d\log GDPreal_{it} + n_i + t_t + v_{it}. \end{aligned} \tag{5}$$

for $i = 1,...,N$, and $t = 2,...,T$, where n_i and v_{it} have the standard error component structure

$$E[n_i] = E[v_{it}] = E[n_i v_{it}] = 0, \tag{6}$$

and where errors are serially uncorrelated:

$$E[v_{it} v_{is}] = 0, \text{ for } s \neq t, \text{ for } i = 1,\ldots,N, \text{and } t = 2,\ldots,T. \tag{7}$$

Where $d\log cpi$ refers to inflation, and $d\log money$, $d\log debt$ and $d\log GDP\,real$ refer to changes in money, public debt, and real GDP, respectively, all in first-difference logarithms; t_t is a set of temporary dichotomous variables to control for possible structural changes in the inflationary process of the countries analyzed, which did not occur in this research, and n_i represents unobserved country-specific effects that seek to capture cross-country heterogeneity in the debt-inflation nexus.

3.2 Data

Data was obtained from the World Bank[5] database and the IMF's International Financial Statistics[6] database. These correspond to a total of 52 countries (20 net-debtor countries of Latin America, including Costa Rica) for the period 1965 to 2014. Classification into developed and developing countries, as well as into creditor or debtor countries was obtained from the *World Economic Outlook 2014* (IMF, 2014).

The variables used in the estimations are described below:

- Gross domestic product at constant 2005 prices in United States dollars, equal to real GDP comparable across countries.
- Historical series for public debt as a percentage of GDP, transformed into real term values by multiplying by the respective real GDP.
- Money and quasi-money (M2) as a percentage of GDP, in the same way as debt, its level is obtained by multiplying by the corresponding GDP.
- Inflation obtained by the GDP deflator, data taken directly from the World Bank for each country.
- Inflation obtained through the log difference of the consumer price index (CPI). This CPI is taken from IMF database for each country and has 2010 as its base year (2010=100).

5 See <http://databank.worldbank.org/data/reports.aspx?source=2&Topic=3>.

6 See <http://data.imf.org/?sk=5804C5E1-0502-4672-BDCD-671BCDC565A9>.

- Country classification: 1, developed countries; 2, net-creditor developing countries; and 3, net-debtor developing countries.

3.3 Estimation Process

The conceptual framework reflected in Equations 1 and 4 posits that the coefficients for debt and money should be positive, and negative for output. In the specifications we assume that β coefficients are constant for each country group.

We also assume that all the explanatory variables for preceding periods represented by X_{it-s} (that is, $d\log cpi_{it-s-1} + d\log money_{it-s} + d\log debt_{it-s} - d\log GDPreal_{it-s}$) are predetermined as follows:

8 $$E\left[X_{it-s}v_{it}\right]=0, \text{ for } s\geq 0,$$

9 $$E\left[X_{it-s}\Delta v_{it}\right]=0, \text{ for } s\geq 1.$$

These two moment conditions allow the use of lagged variables as instruments, once the equation has been first differenced to eliminate specific-country effects (Arellano and Bond, 1991). Given that the variables used in the regressions are not persistent, as shown by the panel data unit root test (Annex 1), we consider instruments in the first differences to be appropriate and that they do not suffer from the weak instrument problem.[7] Hence, we can use the general method of moments estimator (GMM) proposed by Arellano and Bond (1991).

Meanwhile, to test consistency of the estimators for the parameters in Equation 5, besides the first difference GMM estimator, we use a dynamic fixed effect estimator to calculate

[7] Although a weak instrument is exogeneous, it not very important because it is poorly correlated with the endogenous variable it is meant to explain.

an error correction model (ECM) that allows for observing the long-run relations between the variables in a similar way to the GMM estimator, adding estimation of the speed of adjustment from short-run to long-run dynamics. The ECM requires the presence of cointegrating relations between the variables employed, which is verified for the panel of countries studied (Annex 2).

The size of the sample employed means the extent of possible biases in the specification are reduced given that T is higher than 30 for the fixed effects estimator and N is greater than 20 for the GMM estimator (Judson and Owen, 1999). We prefer two-step GMM estimates because the sample size prevents small sample biases. Furthermore, this allows better estimation when regression errors are not distributed identically across countries. The possible existence of serial correlation of errors is handled by using the robust version of each estimator.

Regressions are performed separately for different country groups in order to address a potential problem of slope heterogeneity without sacrificing efficiency gains from panel data. Countries are grouped according to their level of economic development and, among subgroups, sovereign indebtedness as classified by the *World Economic Outlook* 2014 (IMF, 2014) that takes into account balance of payments data[8] from 1972 to 2013. This grouping was considered coherent with the aims of the paper because the criteria is objective and broadly responds to the institutional strength and political credibility of the country sample. Annex 5 shows a detailed list of the countries used and their grouping.

[8] Countries are classified as net debtors when the current account of the balance of payments has accumulated deficits from 1972 to 2013.

4. RESULTS OF THE ESTIMATION

The estimations that include developed countries were not significant and are therefore not shown in the results. The estimations that only include net-debtor developing countries present the best adjustment and significance. The countries included in those estimations are show in Table 1. This group of countries is the most interesting for the study because it allows us to see how inflation reacts in indebted developing countries, such as Costa Rica, to changes in their deficit.

Table 1

COUNTRIES USED IN THE NET-DEBTOR ESTIMATION

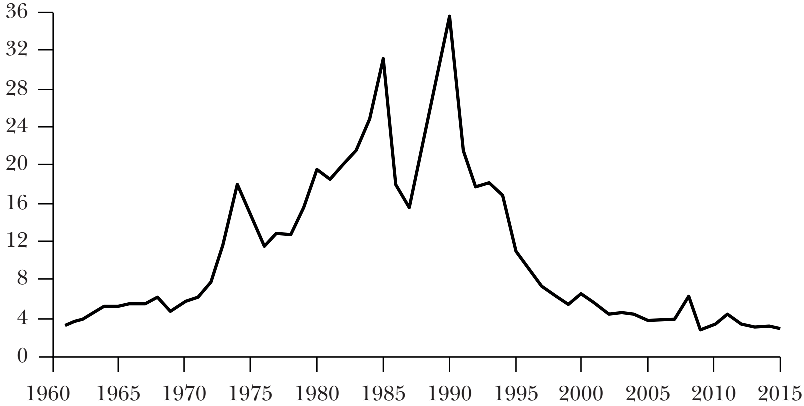
Barbados	Hungary
Brazil	Jamaica
Chile	Mexico
Colombia	Nicaragua
Costa Rica	Panama
Dominican Republic	Paraguay
Ecuador	Peru
El Salvador	Poland
Guatemala	Turkey
Honduras	Uruguay

Source: Own elaboration.

One important aspect to take into account when using time series is the possible existence of structural breaks in the evolution of the variables. Given our use of panel data, we present the average temporal evolution for inflation measured by the CPI (Figure 1), inflation obtained from the GDP deflator (Figure 2), public debt (Figure 3), money, and quasi-money (M2, Figure 4), and economic growth (Figure 5).

Figure 1

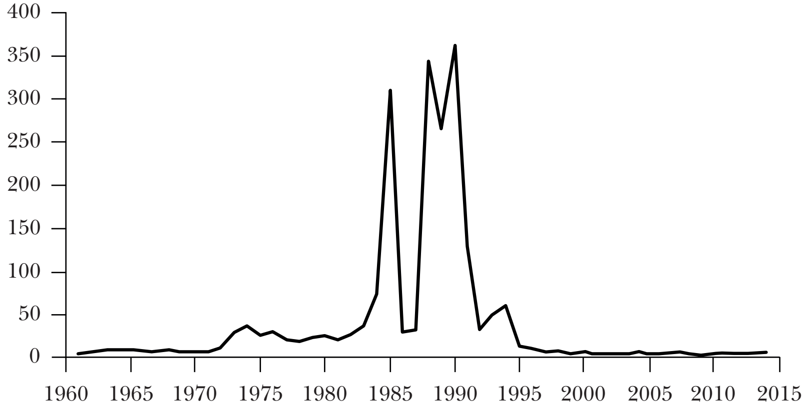
AVERAGE OF CPI INFLATION, 1960-2014
Percentages



Source: Own elaboration.

Figure 2

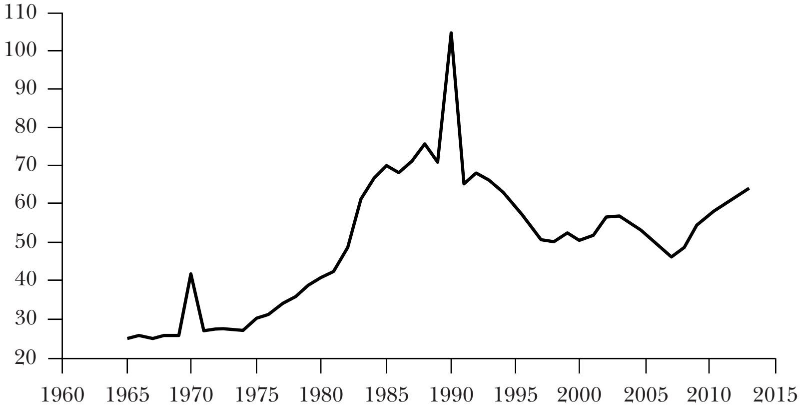
AVERAGE OF INFLATION, BY GDP DEFLATOR, 1960-2014
Percentages



Source: Own elaboration.

Figure 3

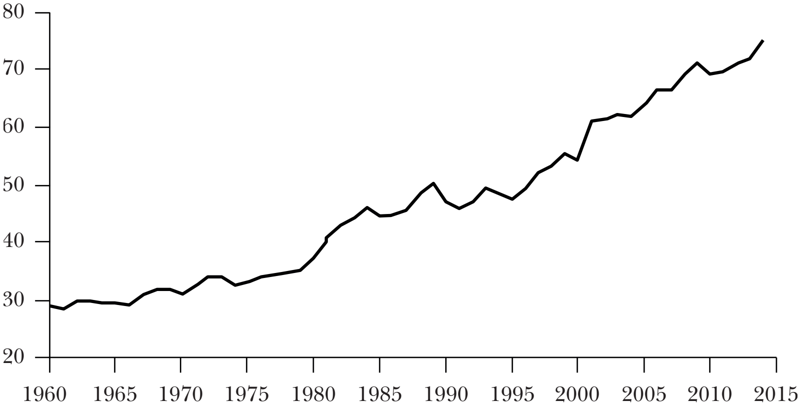
AVERAGE OF PUBLIC DEBT, 1960-2014
As percentage of GDP



Source: Own elaboration.

Figure 4

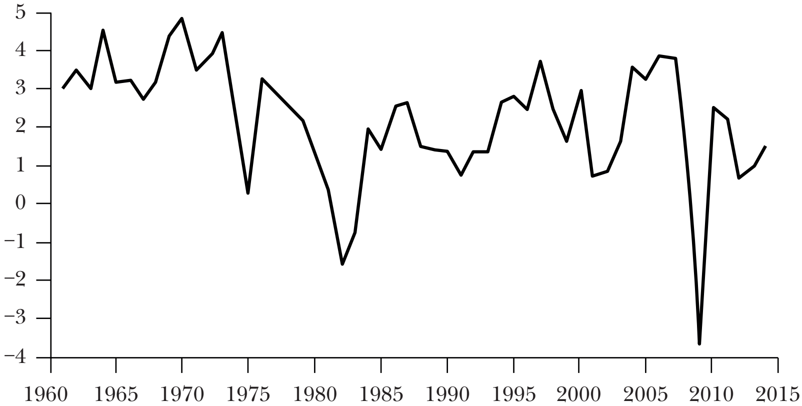
AVERAGE OF MONEY AND QUASI-MONEY (M2), 1960-2014
As percentage of GDP



Source: Own elaboration.

Figure 5

AVERAGE ECONOMIC GROWTH, 1960-2014
Percent



Source: Own elaboration.

As can be seen, the variables employed do not, on average, show evidence for the existence of a structural change in the period studied. The previous figures, together with the descriptive statistics (Annex 3) by year and by country, show that although there are periods of significant hyperinflation, these do not represent a structural change in the inflation data generation process because they are short and followed by downward shifts in inflation, which on average returns to similar levels observed in the period preceding the hyperinflations (Figures 1 and 2). A comparison of this evolution with that of public debt shows that the highest peaks in debt are between the decades of the 1980s and 1990s, corresponding to the periods of greatest hyperinflations. This suggests a direct association between both variables that we seek to test by estimating a panel data model. Although the periods of high debt and hyperinflation are evident, it is important to mention that the use of temporal dichotomous variables to control for said periods would absorb the variability of data in the debt-inflation nexus (stronger during said periods), which could result in the non significance of the relation when it actually might be so. For this reason, temporal dichotomous variables are not used.

Table 2

EXPLANATORY ESTIMATIONS FOR INFLATION
Dependent variable: Inflation as CPI logarithm difference (dlogCPI) or calculated with the GDP deflator

Dependent variable	***Inflation (CPI)***		***Inflation (GDP deflator)***	
Specification	*Dynamic fixed effects panel*	*Arellano-Bond*	*Dynamic fixed effects panel*	*Arellano-Bond*
Speed of adjustment to long run	−0.73[a]	NA	−0.74[a]	NA
Money (M2)	3.65[a] (1.49)	1.95 (5.3)	3.29[a] (1.11)	3.46[c] (1.84)
Total debt	2.9[a] (0.80)	3.56[b] (1.59)	1.76[a] (0.48)	0.95[b] (0.46)
Real GDP	−6.89[a] (2.58)	−5.94 (7.39)	−7.99[a] (1.86)	−6.94[b] (3.49)
Number of observations	424	424	424	424
Number of countries	19	19	19	19

Note: all variables are expressed in logarithms (except inflation). Equations for fixed effects. Standard error in parenthesis. Results in terms of elasticities. Significance at: [a] 1%, [b] 5%, [c] 10 percent.
Source: Own elaboration.

Table 2 summarizes the results of the estimations performed. First, the CPI logarithmic difference is used as a dependent variable that represents inflation. Next, inflation measured by the GDP deflator is employed as a variable to be explained, which is included in levels. Given that explanatory variables are included in log differences, the first estimations produce statistics regarding the price level, while the second ones give semielasticities. In Table 2 the results were transformed in order for them all to be presented as elasticities and allow for their direct comparison. The table also includes the speed of adjustment from short-run dynamics to long-run equilibrium, given by the correction error coefficient of the ECM estimated through the dynamic fixed effects panel. These long-run adjustment values demonstrate that around 74% of an imbalance is corrected during the first year if inflation measured by the GDP deflator is taken as the explanatory variable of inflation, while it would be 73% if the percentage change in the consumer price index is used.

Interpretation of the coefficients gives the relation between the growth rates of explanatory variables and the growth rate of the price level (inflation). An increase of one percentage point (pp) in the growth rate of debt is associated with an increase of between 1 pp and 3.5 pp in the price level over the long-term, it means, if inflation was 3% it would shift to being between 4% and 6.5%. Meanwhile, an increase of 1 pp in the growth rate of money is linked to an increase of between 3.25 pp and 3.65 pp in the growth rate of the price level, again by way of example, a long-term inflation of 3% would shift to being between 6.25% and 6.65% over the long-term. Finally, an increase in the economic growth rate of 1 pp is associated with a decrease of between 6 pp and 8 pp in the inflation rate over the long-term, this means an inflation rate of 10% would shift to being between 4% and 2% in the long run.

These outcomes were in line with other empirical studies on inflation. Many studies report the existence of a positive relation between debt or budget deficits and inflation, mainly in developing countries, but not in developed ones (Feldstein, 1986; Orr et al., 1995; Fischer et al., 2002; Engen and Hubbard, 2004; Catão and Terrones, 2005). In the case of developed economies, numerous studies have even found that there is no link between money and inflation (Dwyer, 1982; Christiano and Fitzgerald, 2003).

Annex 2 shows other estimations performed to provide additional information on the effect of a larger debt on inflation and include the short-term outcomes for error correction estimates (Table A.1) where a relatively greater impact of demand on inflation (GDP) can be seen than that observed in the long-term estimations of Table 2. We also perform the same estimations run for emerging economies, but this time for advanced economies (Annex 2, Table A.2). In this case the amount of money does not have a significant impact on inflation in any of the estimations (dynamic fixed effects or Arellano-Bond), with both measures of inflation. This can be explained by the fact that the monetary channel is less important in such countries because fiscal dominance is much lower. Moreover, the only significant variable, and solely in the dynamic fixed effects estimations, is real GDP, highlighting a greater demand channel effect, as would be expected for advanced economies.

Besides the specifications mentioned above, we attempted to include different types of taxes to observe their effect on inflation. This is important in the context of the need for fiscal reform in Costa Rica. However, none of the tax variables were significant. Likewise, a VAR model was estimated to analyze the transmission channels of the debt, inflation, economic growth, and money supply nexus for Costa Rica, which did not produce positive outcomes either.

5. CONCLUSIONS

This paper provides empirical evidence supporting the hypothesis that, with a net debtor country given, increases in government debt tend to increase inflation, above all in countries with high levels of public debt. The regression results show that an increase in the debt/GDP ratio is significantly and strongly associated with high inflation in indebted developing countries, after controlling for money growth and real output growth. In contrast, this relation is not significant for developed countries.

The outcomes allow for concluding that forward-looking models of inflation are valid for countries such as Costa Rica, in the sense that fiscal policy regimes matter in the debt-inflation nexus. Moreover, certainty regarding cointegrating relations between debt, money, growth, and inflation, even for the panel group of countries, demonstrate that the appropriate conduction of fiscal policy is crucial for macroeconomic stability over the short and long terms.

These findings highlight challenges for price stabilization in highly indebted developing countries because the expansion of public debt affects variables that are sensitive for economic agents' decision making, such as inflation, income, and interest rates. Moreover, despite the important role of monetary policy in managing inflation expectations, fiscal policy could be a dominant factor for the evolution of inflation in highly indebted developing countries. This implies, in general, and for Costa Rica in particular, that price stability achieved through the issuance of central bank instruments could be sustainable only if accompanied by fiscal consolidation and structural reforms that promote monetary policy independence.

Several other aspects are important for future lines of research. First, defining a specification and an appropriate estimation method to study the relation between fiscal variables and inflation in the Costa Rican economy. Second, determining whether the debt-inflation nexus is symmetrical, that is, if increases or decreases in debt have equivalent upward or downward effects on inflation, or if said impact varies in size depending on the direction. Third, investigating the possibility of a non-linear relation between both variables, given that the effect identified in this paper could be much greater for high levels of debt, where governments usually have less credibility and do not have access to credit markets, meaning their only option is to resort to financing from the central bank. Finally, measuring the impact of debt structure, particularly the currency and maturity of sovereign bonds, on inflation dynamics.

ANNEXES

Annex 1. Panel Data Unit Root Test

Panel Data Unit Root Test for Log Public Debt

H_0: All panels contain unit roots	Number of panels	52
H_a: At least one panel is stationary	Average number of periods	42.35
	Augmented Dickey-Fuller	*Phillips-Perron*
Panel statistics	p *value*	p *value*
Inverse χ^2 (102)	0.37	0.89
Inverse normal	0.51	0.97
Inverse logit t(259)	0.47	0.95
Modified inverse χ^2	0.39	0.88

Unit Root Test for Log Money (M2)

H_0: All panels contain unit roots	Number of panels	39
H_a: At least one panel is stationary	Average number of periods	46.87
	Augmented Dickey-Fuller	*Phillips-Perron*
Panel statistics	p *value*	p *value*
Inverse χ^2 (102)	0.21	0.40
Inverse normal	0.07	0.25
Inverse logit t(259)	0.09	0.27
Modified inverse χ^2	0.22	0.42

Unit Root Test for Log GDP		
H_0: All panels contain unit roots	Number of panels	52
H_a: At least one panel is stationary	Average number of periods	49
	Augmented Dickey-Fuller	*Phillips-Perron*
Panel statistics	p *value*	p *value*
Inverse χ^2 (102)	0.99	0.99
Inverse normal	0.99	0.99
Inverse logit t(259)	0.99	0.99
Modified inverse χ^2	0.99	0.99

Unit Root Test for Inflation Measured by GDP Deflator		
H_0: All panels contain unit roots	Number of panels	52
H_a: At least one panel is stationary	Average number of periods	48.81
	Augmented Dickey-Fuller	*Phillips-Perron*
Panel statistics	p *value*	p *value*
Inverse χ^2 (102)	0.00	0.00
Inverse normal	0.00	0.00
Inverse logit t(259)	0.00	0.00
Modified inverse χ^2	0.00	0.00

Unit Root Test for Log CPI		
H_0: All panels contain unit roots	Number of panels	51
H_a: At least one panel is stationary	Average number of periods	49.18
	Augmented Dickey-Fuller	*Phillips-Perron*
Panel statistics	p *value*	p *value*
Inverse χ^2 (102)	1.00	0.16
Inverse normal	1.00	1.00
Inverse logit t(259)	1.00	1.00
Modified inverse χ^2	1.00	0.16

Unit Root Test for Public Debt Log Difference

H_0: All panels contain unit roots	Number of panels	51
H_a: At least one panel is stationary	Average number of periods	41.98
	Augmented Dickey-Fuller	*Phillips-Perron*
Panel statistics	p *value*	p *value*
Inverse χ^2 (102)	0.00	0.00
Inverse normal	0.00	0.00
Inverse logit t(259)	0.00	0.00
Modified inverse χ^2	0.00	0.00

Unit Root Test for Money (M2) Log Difference

H_0: All panels contain unit root	Number of panels	38
H_a: At least one panel is stationary	Average number of periods	47
	Augmented Dickey-Fuller	*Phillips-Perron*
Panel statistics	p *value*	p *value*
Inverse χ^2 (102)	0.00	0.00
Inverse normal	0.00	0.00
Inverse logit t(259)	0.00	0.00
Modified inverse χ^2	0.00	0.00

Unit Root Test for GDP Log Difference

H0: All panels contain unit root	Number of panels	51
Ha: At least one panel is stationary	Average number of periods	48.94
	Augmented Dickey-Fuller	*Phillips-Perron*
Panel statistics	p *value*	p *value*
Inverse χ^2 (102)	0.00	0.00
Inverse normal	0.00	0.00
Inverse logit t(259)	0.00	0.00
Modified inverse χ^2	0.00	0.00

Unit Root Test for CPI Inflation Log Difference		
H0: All panels contain unit root	Number of panels	51
Ha: At least one panel is stationary	Average number of periods	48.18
	Augmented Dickey-Fuller	*Phillips-Perron*
Panel statistics	p *value*	p *value*
Inverse χ^2 (102)	0.00	0.00
Inverse normal	0.00	0.00
Inverse logit t(259)	0.00	0.00
Modified inverse χ^2	0.00	0.00

Source: Own elaboration.

Annex 2. Panel Cointegration Tests and Other Estimations

Table A.1

SHORT-TERM INFLATION ESTIMATIONS
Dependent variable: inflation as the first difference of the log of the CPI or calculated with the GDP deflator

Dependent variable	***Dynamic fixed effects panel: short-term***	
Specification	*CPI*	*Deflator*
Money (M2)	0.33[a] (0.08)	0.34[b] (0.13)
Total debt	0.15[a] (0.05)	0.04 (0.06)
Real GDP	−1.56[a] (0.33)	−1.85[a] (0.52)
Number of observations	424	424
Number of countries	19	19

Note: all variables are expressed in logarithms (except inflation). Fixed effect equations. Standard error in parenthesis. Results in terms of elasticities. Significance at: [a] 1%, [b] 5%, [c] 10 percent.
Source: Own elaboration.

Table A.2

ADVANCED ECONOMIES: EXPLANATORY VARIABLES FOR INFLATION
Dependent variable: inflation as the first difference of the log of the CPI or calculated with the GDP deflator

Dependent variable	*CPI*		*Inflation (GDP Deflator)*	
Specification	*Dynamic fixed effect panel*	*Arellano-Bond*	*Dynamic fixed effect panel*	*Arellano-Bond*
Speed of adjustment	−0.39[a]	na	−0.30[a]	na
Money (M2)	0.266 (0.39)	0.02 (0.03)	0.262 (0.74)	0.49 (2.06)
Total debt	0.44[b] (0.19)	0.09 (0.19)	0.47 (0.38)	1.87 (2.36)
Real GDP	−1.66[b] (0.74)	−0.49 (0.67)	−2.69[b] (1.35)	−0.53 (0.92)
Number of observations	331	331	331	331
Number of countries	16	16	16	16

Note: all variables are expressed in logarithms (except inflation). Fixed effect equations. Standard error in parenthesis. Results in terms of elasticities. Significance at: [a] 1%, [b] 5%, [c] 10 percent.
Source: Own elaboration.

Pedroni Panel Cointegration Test

H_0: No-cointegration		
H_a: Common AR coefficients (within dimensions)	*Non-weighted*	*Weighted*
Panel statistics	*p* value	*p* value
Panel v statistic	0.00	0.51
Panel ρ statistic	0.00	0.00
Panel PP statistic	0.00	0.00
Panel ADF statistic	0.00	0.00
H_a: Individual AR coefficients (within dimensions)		p *value*
Group ρ statistic		0.00
Group PP statistic		0.00
Group ADF statistic		0.00

Source: Own elaboration.

Annex 3. Descriptive Statistics of Inflation, Public Debt, Money (M2) and Economic Growth by Country and by Year

Descriptive Statistics of Public Debt (as a percentage of GDP) by Country

Country	Average	Mean	Max.	Min.	Standard deviation	Obs.
Argentina	40.8	36.3	137.7	9.3	28.0	43
Australia	23.4	22.7	41.2	9.7	7.6	48
Austria	49.0	56.2	82.3	12.8	22.5	48
Barbados	46.4	46.4	96.3	15.8	17.2	41
Belgium	92.6	100.3	138.4	38.8	32.9	48
Bolivia	83.9	79.0	205.2	32.5	39.0	44
Brazil	58.7	62.6	102.9	29.9	17.5	36
Canada	69.9	71.1	100.8	42.8	17.0	49
Chile	44.6	28.9	165.5	3.9	44.5	44
Colombia	26.0	28.3	44.7	9.2	10.4	49
Costa Rica	43.0	33.3	110.3	18.1	26.3	49
Czech Republic	25.6	27.8	45.1	11.6	10.5	21
Denmark	39.1	45.0	78.6	4.3	23.1	49
Dominican Republic	30.1	25.6	60.7	12.7	13.6	44
Ecuador	58.2	29.7	661.2	14.6	92.3	49
El Salvador	39.2	32.4	108.3	10.2	23.9	49
Estonia	6.1	5.7	9.9	3.7	1.8	19
Finland	26.6	17.0	57.6	1.7	19.5	47
France	41.4	34.2	92.3	14.4	23.9	49
Germany	44.0	40.4	80.6	17.6	19.5	49
Greece	65.1	56.2	175.0	0.0	45.8	49
Guatemala	25.3	21.5	55.6	10.1	13.4	48
Honduras	52.9	48.1	243.4	6.5	41.4	49
Hungary	80.3	78.3	127.6	51.8	21.5	30
Iceland	40.2	34.3	95.1	11.8	22.1	42
Ireland	60.3	53.0	120.2	23.6	25.4	49
Israel	113.0	98.4	284.0	62.1	49.5	40

Descriptive Statistics of Public Debt (as a Percentage of GDP) by Country

Country	*Average*	*Mean*	*Max.*	*Min.*	*Standard deviation*	*Obs.*
Italy	82.6	93.3	128.5	28.4	29.8	49
Jamaica	90.1	92.5	181.3	14.2	48.8	48
Japan	93.6	71.2	242.6	5.2	71.6	49
Korea	17.0	16.8	34.5	2.3	8.3	49
Luxemburg	8.6	7.1	23.0	2.2	4.7	42
Mexico	39.4	41.8	78.1	5.7	16.5	47
Netherlands	57.8	55.6	78.5	37.8	12.5	49
New Zealand	43.4	46.4	76.0	14.6	16.3	49
Nicaragua	177.3	92.6	2,092.9	0.7	315.8	49
Norway	34.3	32.5	52.6	22.3	9.2	48
Panama	62.1	64.8	115.8	17.8	26.0	49
Paraguay	28.0	22.2	67.0	13.0	13.6	44
Peru	37.9	37.1	63.4	19.0	11.6	44
Poland	53.9	49.6	90.1	36.8	14.5	28
Portugal	48.9	52.7	129.7	13.5	27.1	49
Slovakia	37.8	38.6	54.6	21.4	9.5	22
Slovenia	29.5	26.3	70.5	16.8	13.1	21
Spain	37.4	40.0	92.1	7.3	22.0	49
Sweden	47.0	47.7	70.9	16.1	16.8	47
Switzerland	43.2	45.6	67.0	7.0	14.7	48
Turkey	37.5	34.6	77.9	19.0	12.8	49
United Kingdom	54.2	48.7	94.6	31.0	17.5	49
United States	56.1	57.4	104.8	32.2	19.1	49
Uruguay	52.0	42.3	111.5	16.6	27.7	44
Venezuela	30.7	31.6	71.9	4.6	19.6	47
All	51.5	42.0	2,092.9	0.0	60.8	2,300

Descriptive Statistics of Public Debt (as a Percentage of GDP) by Year

Year	*Average*	*Mean*	*Max.*	*Min.*	*Standard deviation*	*Obs.*
1965	24.8	18.4	94.6	5.2	21.3	28
1966	25.8	19.1	91.9	4.4	21.0	31
1967	24.9	19.5	89.1	3.7	19.8	34
1968	25.4	19.9	88.5	2.7	19.6	34
1969	25.4	22.0	82.8	0.7	18.9	33
1970	41.9	22.3	661.2	2.3	99.3	42
1971	26.8	22.4	65.7	4.6	15.6	42
1972	27.3	23.4	77.7	2.2	17.5	45
1973	27.1	21.6	100.9	2.5	19.8	45
1974	26.9	22.6	79.8	1.7	17.5	44
1975	30.1	24.1	108.3	2.0	20.9	44
1976	31.1	26.3	97.4	0.0	20.7	44
1977	33.8	28.2	142.0	0.0	24.8	45
1978	35.8	31.6	133.6	0.0	24.4	46
1979	38.7	32.3	155.5	7.1	28.6	46
1980	40.6	30.9	154.3	6.4	31.6	46
1981	42.4	35.4	149.1	6.7	30.2	44
1982	48.4	38.5	159.1	6.9	32.6	46
1983	61.1	48.2	260.5	7.4	47.8	47
1984	66.1	53.6	284.0	7.7	50.8	47
1985	70.2	56.5	218.0	6.3	49.5	47
1986	68.0	56.3	169.6	7.9	39.7	48
1987	71.2	59.9	266.6	6.7	45.3	48
1988	75.9	59.3	629.2	5.1	87.5	48
1989	70.8	58.9	477.0	4.0	67.7	47

Year	Average	Mean	Max.	Min.	Standard deviation	Obs.
1990	105.2	55.9	2,092.9	4.7	295.7	48
1991	65.1	51.8	333.7	4.0	50.8	48
1992	68.1	49.5	448.6	4.8	64.4	48
1993	66.1	52.6	445.9	6.0	63.3	49
1994	63.2	50.1	446.6	5.5	62.6	50
1995	59.3	54.1	362.7	8.9	51.7	51
1996	55.2	55.2	222.4	7.4	36.8	52
1997	50.4	49.4	123.6	6.1	28.5	52
1998	50.1	45.5	121.6	5.5	28.3	52
1999	52.5	47.1	135.6	6.0	29.4	51
2000	50.6	46.3	143.8	5.1	28.5	52
2001	51.6	48.4	153.6	4.8	29.1	52
2002	56.7	50.9	164.0	5.7	33.8	52
2003	57.0	48.5	169.6	5.6	32.8	52
2004	54.9	45.9	180.7	5.1	32.4	52
2005	52.4	46.3	186.4	4.5	32.1	52
2006	49.4	42.8	186.0	4.4	31.8	52
2007	46.1	38.1	183.0	3.7	31.6	52
2008	48.5	41.2	191.8	4.5	33.4	52
2009	54.6	45.8	210.2	5.8	37.0	52
2010	57.3	43.5	215.8	6.5	38.7	52
2011	59.7	46.3	229.7	5.9	42.1	52
2012	62.1	50.2	236.6	9.5	42.8	52
2013	64.3	53.3	242.6	9.9	43.8	52
All	51.5	42.0	2,092.9	0.0	60.8	2,300

Descriptive Statistics of Money (M2, as a Percentage of GDP) by Country

Country	*Average*	*Mean*	*Max.*	*Min.*	*Standard deviation*	*Obs.*
Argentina	22.0	22.4	31.8	10.6	4.7	55
Australia	59.2	48.6	109.5	37.7	20.6	55
Barbados	67.0	55.4	118.9	37.8	27.4	30
Bolivia	32.9	21.0	81.2	6.4	22.0	55
Brazil	36.8	24.8	111.3	10.1	24.3	55
Canada	72.8	65.1	158.1	36.2	33.9	49
Chile	44.9	37.6	96.2	11.2	27.3	54
Colombia	28.9	28.6	46.8	19.6	6.7	53
Costa Rica	33.6	31.9	56.9	14.6	12.6	55
Czech Republic	63.3	61.8	78.2	53.4	7.5	22
Denmark	51.7	50.5	70.1	40.0	8.0	55
Dominican Republic	26.9	26.9	50.2	14.4	7.0	55
Ecuador	17.5	15.7	33.3	7.8	5.8	55
El Salvador	36.0	37.4	52.8	20.0	9.4	50
Estonia	39.7	31.6	62.4	16.2	17.8	16
Guatemala	26.7	23.0	47.2	12.9	10.8	55
Honduras	33.1	30.9	56.8	14.9	13.0	55
Hungary	51.7	49.9	63.3	44.1	6.5	24
Iceland	43.8	37.0	102.8	19.6	23.1	55
Israel	64.7	72.8	133.4	21.7	25.5	55
Jamaica	46.9	48.6	73.1	17.7	14.1	55

Country	*Average*	*Mean*	*Max.*	*Min.*	*Standard deviation*	*Obs.*
Japan	163.4	181.1	251.3	48.5	58.1	55
Korea	54.2	33.3	139.9	8.9	42.5	55
Mexico	27.9	27.1	38.7	11.0	4.5	55
New Zealand	49.7	30.6	93.5	19.8	27.1	50
Nicaragua	28.9	28.0	69.9	12.1	14.7	55
Norway	51.8	51.4	59.4	47.7	3.2	47
Panama	50.0	42.0	87.2	16.2	23.5	55
Paraguay	24.7	24.2	50.6	9.5	9.0	55
Peru	25.8	24.2	43.1	16.6	7.1	55
Poland	43.2	42.3	61.6	30.4	9.7	25
Slovakia	59.8	59.8	65.1	55.3	2.7	16
Sweden	51.4	51.1	67.1	38.2	7.8	55
Switzerland	118.8	110.7	188.6	90.6	24.3	45
Turkey	30.4	25.6	60.6	14.6	12.5	55
United Kingdom	72.7	56.1	170.2	30.5	44.4	55
United States	70.7	69.8	90.4	59.5	7.4	54
Uruguay	38.9	39.5	63.9	14.5	11.9	55
Venezuela	28.1	28.4	52.9	16.4	8.7	54
All	47.5	37.9	251.3	6.4	34.7	1,909

Descriptive Statistics of Money (M2, as a Percentage of GDP) by Year

Year	*Average*	*Mean*	*Max.*	*Min.*	*Standard deviation*	*Obs.*
1965	29.8	20.8	96.0	7.8	20.3	33
1966	29.7	20.6	95.8	9.2	20.2	33
1967	31.1	21.2	95.6	10.2	20.2	33
1968	31.8	20.9	99.7	12.6	20.6	33
1969	32.1	21.1	101.2	11.8	20.2	33
1970	31.3	22.7	103.2	11.3	19.0	31
1971	32.5	24.4	116.5	13.7	20.3	32
1972	34.1	26.0	127.1	14.7	21.1	32
1973	34.2	26.7	124.2	14.5	20.9	32
1974	32.9	26.4	118.7	13.0	20.4	32
1975	33.5	28.6	125.7	12.9	21.2	32
1976	34.3	27.3	129.3	15.2	21.1	32
1977	34.5	28.7	131.9	14.5	21.7	32
1978	34.9	28.2	137.1	14.5	22.4	32
1979	35.5	28.8	140.6	15.1	22.9	32
1980	37.3	29.6	142.2	11.9	24.3	34
1981	40.9	32.7	147.8	11.3	26.6	34
1982	43.1	35.3	153.9	10.1	28.2	34
1983	44.2	35.6	160.5	10.2	29.9	34
1984	46.1	37.6	162.9	11.5	32.4	34
1985	44.8	39.5	164.9	12.1	30.5	34
1986	44.8	39.6	172.1	10.7	30.7	33
1987	45.8	35.2	181.1	13.8	32.0	34
1988	48.6	36.8	183.7	11.0	33.1	34
1989	50.3	38.4	189.3	10.2	35.4	33
1990	47.3	34.1	187.4	11.5	32.4	35

Year	Average	Mean	Max.	Min.	Standard deviation	Obs.
1991	46.1	35.6	186.5	10.6	31.8	36
1992	47.2	37.5	188.1	13.0	31.3	36
1993	49.6	40.4	195.2	18.2	32.3	38
1994	48.5	42.9	201.4	13.9	32.3	38
1995	47.6	36.9	207.2	15.7	33.6	39
1996	49.5	40.1	210.7	18.4	34.4	39
1997	52.2	42.3	218.2	14.7	36.1	39
1998	53.2	44.2	229.8	14.6	38.1	39
1999	55.6	47.8	239.7	13.1	40.0	39
2000	54.5	45.0	240.6	17.4	38.7	39
2001	61.2	48.4	200.8	19.7	37.5	39
2002	61.5	47.0	205.2	17.4	38.6	39
2003	62.1	50.2	206.5	18.2	38.1	39
2004	62.0	49.4	205.7	21.1	38.2	39
2005	64.0	52.7	206.6	22.3	38.9	39
2006	66.5	55.1	204.0	22.3	39.8	39
2007	66.7	56.9	202.8	24.2	38.3	38
2008	69.2	56.6	209.1	21.0	41.3	38
2009	71.1	60.2	227.0	22.4	43.3	36
2010	69.3	58.2	226.1	23.2	42.7	35
2011	69.5	56.3	238.0	22.9	45.0	33
2012	71.0	56.6	241.3	25.8	46.0	33
2013	71.9	58.6	247.8	26.7	47.1	32
2014	75.1	61.0	251.3	26.6	47.4	32
All	47.5	37.9	251.3	6.4	34.7	1,909

Descriptive Statistics for CPI Inflation (as Percentage) by Country

Country	*Average*	*Mean*	*Max.*	*Min.*	*Standard deviation*	*Obs.*
Australia	4.8	3.3	14.1	−0.3	3.6	55
Austria	3.3	3.0	9.1	0.5	1.9	55
Barbados	6.0	4.9	32.9	−1.3	5.8	49
Belgium	3.5	2.7	12.0	−0.1	2.7	55
Bolivia	28.4	7.6	477.5	−0.7	74.9	55
Brazil	76.2	8.6	341.7	3.1	103.5	35
Canada	3.8	2.7	11.7	0.2	2.9	55
Chile	3.0	3.1	4.3	1.4	1.2	6
Colombia	13.9	15.5	29.1	2.0	8.0	55
Costa Rica	11.3	9.6	64.2	−0.7	10.3	55
Czech Republic	3.8	2.6	10.1	0.1	3.2	22
Denmark	4.7	3.4	14.2	0.5	3.3	55
Dominican Republic	10.2	7.4	41.5	−4.0	11.0	55
Ecuador	16.8	11.0	67.3	2.3	15.7	55
El Salvador	7.3	4.5	27.7	−2.7	7.2	55
Estonia	9.8	4.0	64.1	−0.5	15.0	23
Finland	4.8	3.9	16.4	−0.2	4.1	55
France	4.2	2.7	12.8	0.0	3.5	55
Germany	2.0	2.3	6.8	−35.4	5.4	55
Greece	8.2	4.6	23.8	−1.8	7.2	55
Guatemala	7.7	6.6	34.5	−0.8	7.3	55
Honduras	8.0	6.5	29.2	1.1	6.2	55
Hungary	9.1	6.6	29.4	−0.2	7.3	43
Iceland	14.4	9.3	61.1	1.5	14.1	55
Ireland	5.5	3.9	19.0	−4.6	5.2	55
Israel	21.3	8.6	155.6	−0.6	33.3	55

Country	Average	Mean	Max.	Min.	Standard deviation	Obs.
Italy	5.9	4.3	19.3	0.0	5.1	55
Jamaica	13.0	9.2	57.3	1.4	10.4	55
Japan	3.1	2.0	20.8	−1.4	4.0	55
Korea	7.1	4.6	25.2	0.7	6.1	49
Luxemburg	3.4	2.8	10.2	−0.1	2.5	55
Mexico	16.6	6.7	84.1	0.6	19.9	55
Netherlands	3.4	2.6	9.7	−0.7	2.4	55
New Zealand	5.6	3.4	15.8	0.2	4.8	55
Nicaragua	7.3	6.9	18.1	3.6	3.5	16
Norway	4.5	3.4	12.8	0.5	3.1	55
Panama	2.8	1.6	15.1	−0.1	3.1	55
Paraguay	10.4	8.8	31.7	−0.9	7.8	55
Peru	39.6	9.1	432.8	0.2	79.9	55
Poland	19.0	6.8	188.0	−1.0	34.3	45
Portugal	8.1	4.9	25.3	−0.8	7.4	55
Slovakia	5.1	4.4	12.6	−0.3	3.6	22
Slovenia	6.4	5.4	28.4	−0.5	6.6	23
Spain	6.6	5.1	21.9	−0.5	5.1	55
Sweden	4.4	3.4	12.8	−0.5	3.6	55
Switzerland	2.6	1.9	9.3	−1.2	2.3	55
Turkey	26.4	17.6	74.3	0.4	21.5	55
United Kingdom	2.6	2.3	7.3	0.1	1.7	27
United States	3.8	3.1	12.7	−0.4	2.7	55
Uruguay	31.8	29.3	81.2	4.3	22.5	55
Venezuela	36.2	24.8	79.6	19.1	21.5	7
All	10.7	4.7	477.5	−35.4	25.7	2,457

Descriptive Statistics for CPI Inflation (as Percentage) by Year

Year	*Average*	*Mean*	*Max.*	*Min.*	*Standard deviation*	*Obs.*
1965	5.1	3.6	44.8	−1.9	7.3	38
1966	5.5	3.8	55.1	−1.2	8.8	38
1967	5.4	3.3	63.8	0.5	9.8	40
1968	6.2	3.8	81.2	0.0	12.6	40
1969	4.7	3.2	20.0	−0.2	4.2	40
1970	5.7	4.8	15.1	−0.9	3.4	40
1971	6.2	5.7	21.5	−0.5	4.0	41
1972	7.7	6.3	56.8	−0.1	8.4	41
1973	11.7	9.5	67.8	2.4	10.2	42
1974	18.0	15.2	57.2	1.8	10.9	42
1975	14.8	13.0	59.6	2.2	10.1	42
1976	11.5	9.3	41.0	1.7	8.0	42
1977	13.0	10.6	45.9	1.3	9.4	42
1978	12.7	9.0	45.6	1.1	11.3	42
1979	15.5	10.7	57.8	3.6	13.8	42
1980	19.6	14.0	83.7	3.9	17.3	42
1981	18.5	12.8	77.4	4.4	16.1	43
1982	20.1	10.6	80.4	0.3	22.1	43
1983	21.7	9.0	132.3	1.9	28.9	43
1984	25.0	8.2	262.6	1.6	47.4	43
1985	31.2	8.5	477.5	1.0	75.9	43
1986	18.0	7.4	132.5	−0.1	26.7	43
1987	15.5	8.1	118.9	−0.7	23.4	43
1988	22.9	6.9	203.7	0.4	43.4	43
1989	28.9	6.9	355.5	0.2	67.1	44
1990	35.7	9.1	432.8	0.8	83.8	44

Year	Average	Mean	Max.	Min.	Standard deviation	Obs.
1991	21.7	8.1	167.3	−35.4	36.5	44
1992	17.8	5.4	235.3	1.0	37.2	44
1993	18.2	4.6	301.0	0.5	45.1	46
1994	16.9	5.6	308.0	0.1	45.0	48
1995	11.1	4.8	63.2	−0.1	13.3	48
1996	9.2	4.4	59.0	0.1	10.8	48
1997	7.4	4.4	61.9	0.3	10.1	48
1998	6.5	2.6	61.3	−1.3	10.0	48
1999	5.4	2.3	50.0	−0.3	9.3	48
2000	6.6	3.4	67.3	−0.7	10.9	49
2001	5.6	4.0	43.4	−0.8	7.3	49
2002	4.4	3.0	37.1	−1.3	5.7	49
2003	4.6	2.6	24.3	0.1	5.4	49
2004	4.4	2.8	41.5	−0.4	6.2	49
2005	3.8	2.7	14.2	−0.3	3.0	49
2006	3.7	3.1	10.9	0.2	2.5	49
2007	3.9	2.8	10.6	0.1	2.6	49
2008	6.2	4.5	19.9	1.4	4.0	49
2009	2.7	1.8	24.0	−4.6	4.2	50
2010	3.4	2.4	24.8	−1.0	3.8	51
2011	4.5	3.5	23.2	−0.3	3.4	51
2012	3.6	3.0	19.1	−0.7	2.9	51
2013	3.1	1.8	34.1	−0.9	4.9	51
2014	3.1	1.6	48.3	−1.3	6.9	51
2015	2.9	0.6	79.6	−1.8	11.2	51
All	10.7	4.7	477.5	−35.4	25.7	2,457

Descriptive Statistics for GDP Deflator Inflation (as Percentage) by Country

Country	*Average*	*Mean*	*Max.*	*Min.*	*Standard deviation*	*Obs.*
Argentina	178.5	27.0	3,058.0	−2.0	503.4	54
Australia	5.2	5.0	16.0	0.0	3.9	54
Austria	3.3	3.0	10.0	0.0	2.1	54
Barbados	5.9	4.5	31.0	−5.0	7.6	54
Belgium	3.6	3.0	13.0	0.0	2.6	54
Bolivia	277.1	8.0	12,339.0	−5.0	1,684.3	54
Brazil	231.6	32.5	2,700.0	5.0	565.6	54
Canada	4.1	3.0	15.0	−2.0	3.3	54
Chile	49.9	13.5	665.0	0.0	114.8	54
Colombia	16.4	16.5	45.0	2.0	9.7	54
Costa Rica	13.8	11.0	84.0	−1.0	14.4	54
Czech Republic	6.3	3.0	36.0	−1.0	8.2	24
Denmark	4.9	4.0	13.0	0.0	3.4	54
Dominican Republic	12.0	6.0	103.0	−2.0	17.9	54
Ecuador	6.1	5.0	97.0	−26.0	17.2	54
El Salvador	4.8	4.0	18.0	−1.0	4.8	49
Estonia	6.3	5.0	24.0	0.0	5.1	19
Finland	5.3	4.5	22.0	0.0	4.4	54
France	4.4	3.0	14.0	0.0	3.8	54
Germany	2.6	2.0	8.0	0.0	2.0	44
Greece	9.2	5.0	27.0	−3.0	8.1	54
Guatemala	8.1	6.5	41.0	−4.0	9.0	54
Honduras	8.7	6.0	31.0	−3.0	7.4	54
Hungary	9.8	5.0	27.0	2.0	8.0	23
Iceland	17.1	11.0	77.0	0.0	17.0	54
Ireland	6.2	5.0	21.0	−4.0	6.0	44
Israel	33.4	9.0	391.0	−2.0	69.0	54

Country	Average	Mean	Max.	Min.	Standard deviation	Obs.
Italy	6.8	4.5	21.0	0.0	5.8	54
Jamaica	16.7	12.0	60.0	−5.0	12.8	42
Japan	2.8	2.0	23.0	−2.0	5.0	54
Korea	9.8	6.0	33.0	−1.0	8.5	54
Luxemburg	4.1	4.0	20.0	−4.0	4.3	53
Mexico	21.6	9.5	140.0	1.0	28.4	54
Netherlands	3.5	2.0	13.0	−1.0	2.9	54
New Zealand	5.3	3.0	17.0	0.0	5.2	36
Nicaragua	545.0	10.0	13,612.0	−1.0	2,116.8	54
Norway	5.2	5.0	15.0	−5.0	4.0	54
Panama	3.9	2.0	34.0	−1.0	5.4	54
Paraguay	11.7	10.0	38.0	−2.0	9.4	54
Peru	205.5	10.0	6,261.0	0.0	912.4	54
Poland	11.8	4.0	55.0	1.0	15.0	24
Portugal	8.3	4.0	26.0	0.0	7.9	54
Slovakia	4.6	4.0	16.0	−1.0	4.3	22
Slovenia	4.3	4.0	11.0	−1.0	3.4	19
Spain	7.1	6.0	23.0	0.0	5.4	54
Sweden	4.9	4.0	15.0	0.0	3.6	54
Switzerland	1.8	1.0	7.0	0.0	1.9	33
Turkey	34.6	23.5	138.0	2.0	31.8	54
United Kingdom	5.6	4.0	26.0	1.0	5.0	54
United States	3.4	3.0	9.0	1.0	2.3	54
Uruguay	41.3	30.0	192.0	1.0	36.9	54
Venezuela	22.5	16.5	116.0	0.0	22.4	54
All	40.3	5.0	13,612.0	−26.0	438.2	2,538

Descriptive Statistics for GDP Deflator Inflation (as Percentage) by Year

Year	*Average*	*Mean*	*Max.*	*Min.*	*Standard deviation*	*Obs.*
1965	10.1	4.0	97.0	−5.0	19.5	40
1966	8.2	4.0	72.0	−2.0	12.9	41
1967	7.0	3.0	79.0	−2.0	13.4	42
1968	8.2	4.0	116.0	−5.0	18.5	42
1969	7.1	5.0	40.0	0.0	7.7	42
1970	7.4	5.0	41.0	−14.0	8.2	42
1971	7.5	6.5	32.0	−9.0	7.5	44
1972	12.2	7.0	86.0	−5.0	17.4	44
1973	28.3	13.0	414.0	5.0	66.1	44
1974	36.4	19.5	665.0	6.0	97.7	44
1975	27.5	14.0	335.0	−1.0	56.0	44
1976	29.6	12.0	438.0	3.0	73.0	44
1977	21.6	12.5	159.0	1.0	27.9	44
1978	19.3	9.0	161.0	1.0	27.1	45
1979	23.6	14.0	147.0	3.0	27.9	45
1980	26.2	18.0	135.0	4.0	27.0	45
1981	22.3	12.5	126.0	3.0	27.4	46
1982	27.0	10.0	208.0	−9.0	43.3	46
1983	36.1	9.0	382.0	−14.0	72.4	46
1984	72.2	8.0	1,443.0	−4.0	232.8	46
1985	312.4	7.5	12,339.0	−2.0	1,815.7	46
1986	29.4	6.5	281.0	−14.0	56.7	46
1987	32.8	7.0	523.0	−9.0	84.4	46
1988	344.9	7.0	13,612.0	−12.0	2,004.2	46
1989	264.3	7.5	4,709.0	−1.0	899.6	46

Year	*Average*	*Mean*	*Max.*	*Min.*	*Standard deviation*	*Obs.*
1990	364.0	9.5	6,261.0	−1.0	1,246.5	46
1991	129.8	8.0	4,524.0	0.0	652.8	48
1992	32.5	6.0	968.0	−1.0	137.5	49
1993	51.8	5.5	2,001.0	−1.0	281.7	50
1994	61.0	7.5	2,303.0	0.0	324.5	50
1995	12.9	5.0	94.0	−1.0	19.4	50
1996	11.1	4.5	116.0	−1.0	19.3	52
1997	8.0	4.0	81.0	−2.0	12.8	52
1998	8.1	5.0	138.0	−4.0	19.2	52
1999	4.8	3.0	54.0	−26.0	9.6	52
2000	6.5	3.5	49.0	−8.0	9.5	52
2001	5.4	4.0	53.0	−4.0	8.1	52
2002	5.6	3.0	37.0	−2.0	7.9	51
2003	5.5	3.0	35.0	−2.0	7.4	51
2004	6.2	3.0	45.0	−1.0	8.4	51
2005	4.5	3.0	30.0	−1.0	4.7	51
2006	5.0	4.0	18.0	−1.0	4.1	51
2007	5.1	4.0	18.0	−1.0	3.8	52
2008	6.0	4.0	30.0	−2.0	5.7	52
2009	2.8	2.0	12.0	−5.0	3.4	52
2010	4.1	3.0	46.0	−4.0	7.0	52
2011	4.4	3.0	28.0	−3.0	5.2	52
2012	3.2	2.0	19.0	−1.0	3.7	52
2013	3.5	2.0	36.0	−2.0	5.9	52
2014	4.0	2.0	49.0	−3.0	8.0	48
All	40.3	5.0	13,612.0	−26.0	438.2	2,538

Descriptive Statistics for Economic Growth (Percentage) by Country

Country	*Average*	*Mean*	*Max.*	*Min.*	*Standard deviation*	*Obs.*
Argentina	1.4	2.1	10.6	−12.5	5.6	54
Australia	1.9	2.0	5.0	−3.5	1.7	54
Austria	2.5	2.4	8.6	−4.1	2.1	54
Barbados	1.7	1.5	10.5	−17.1	4.9	54
Belgium	2.3	2.2	7.1	−3.5	2.1	54
Bolivia	0.8	2.1	5.2	−15.3	3.6	54
Brazil	2.3	2.1	10.7	−6.8	3.6	54
Canada	2.0	2.1	6.0	−4.3	2.1	54
Chile	2.6	3.3	9.7	−13.6	4.5	54
Colombia	2.1	2.4	5.8	−6.0	2.0	54
Costa Rica	2.2	2.8	6.8	−10.4	3.2	54
Czech Republic	1.5	1.9	6.4	−12.1	4.0	24
Denmark	1.9	2.0	8.1	−5.8	2.3	54
Dominican Republic	2.9	3.1	13.9	−16.5	5.0	54
Ecuador	1.6	1.7	10.3	−6.9	2.9	54
El Salvador	0.9	1.5	5.7	−14.2	3.9	49
Estonia	4.8	6.9	12.2	−15.7	6.4	19
Finland	2.5	2.5	9.2	−9.1	3.2	54
France	2.2	2.0	6.1	−3.5	1.9	54
Germany	1.9	1.9	5.3	−5.5	2.0	44
Greece	2.3	2.4	10.2	−9.0	4.4	54
Guatemala	1.3	1.5	6.5	−6.0	2.3	54
Honduras	1.1	1.5	6.9	−5.1	3.0	54
Hungary	2.0	3.0	4.9	−6.6	2.8	23
Iceland	2.5	2.8	11.5	−7.0	3.9	54
Ireland	3.1	2.8	9.2	−7.6	3.4	44

Country	*Average*	*Mean*	*Max.*	*Min.*	*Standard deviation*	*Obs.*
Israel	2.9	2.5	17.1	−2.6	3.7	54
Italy	2.1	1.9	8.1	−6.1	2.8	54
Japan	3.1	2.3	11.8	−5.7	3.5	54
Korea	5.7	5.9	12.0	−6.6	3.7	54
Luxemburg	2.5	2.5	9.1	−7.9	3.4	53
Mexico	1.8	2.1	8.1	−7.9	3.2	54
Netherlands	2.2	2.1	11.5	−3.9	2.5	54
New Zealand	1.4	1.6	5.1	−5.9	2.1	36
Nicaragua	0.2	1.8	10.2	−33.7	6.5	54
Norway	2.5	2.7	5.6	−2.9	1.9	54
Panama	2.8	3.3	9.6	−16.5	4.3	54
Paraguay	2.3	2.4	11.6	−6.1	3.8	54
Peru	1.5	2.2	9.8	−15.3	4.9	54
Poland	3.6	3.9	7.0	−7.6	2.9	24
Portugal	3.0	3.0	16.2	−8.2	4.0	54
Slovakia	3.8	4.6	10.1	−5.6	3.2	22
Slovenia	2.3	3.4	6.2	−9.0	3.5	19
Spain	2.6	2.3	10.3	−4.5	2.9	54
Sweden	2.1	2.1	9.7	−6.2	2.5	54
Switzerland	1.0	1.0	3.4	−3.4	1.7	33
Turkey	2.5	3.0	8.3	−7.3	3.8	54
United Kingdom	2.1	2.2	9.5	−5.2	2.3	54
United States	2.0	2.1	6.1	−3.7	2.0	54
Uruguay	1.7	1.9	7.8	−11.5	4.3	54
Venezuela	0.0	−0.3	15.0	−11.5	5.1	54
All	2.1	2.3	17.1	−33.7	3.6	2,496

Descriptive Statistics for Economic Growth (Percentage) by Year

Year	Average	Mean	Max.	Min.	Standard deviation	Obs.
1961	3.0	3.3	10.3	−5.7	3.7	40
1962	3.5	3.4	12.4	−2.8	3.1	40
1963	3.1	3.3	9.3	−7.0	3.4	40
1964	4.5	4.5	10.0	0.2	2.4	40
1965	3.2	3.6	10.5	−16.5	4.2	40
1966	3.2	2.9	9.5	−3.3	3.2	41
1967	2.8	2.5	9.9	−4.7	2.8	41
1968	3.2	3.7	13.0	−15.3	4.7	41
1969	4.4	4.2	10.9	−2.6	3.3	41
1970	4.8	4.1	16.2	−2.9	4.3	41
1971	3.5	3.0	11.5	−1.9	2.7	43
1972	3.9	4.1	10.3	−2.6	2.9	43
1973	4.5	4.5	12.0	−6.8	3.4	43
1974	2.2	2.8	10.2	−7.0	3.4	43
1975	0.3	−0.1	7.7	−13.6	3.9	43
1976	3.3	3.6	11.0	−3.6	2.7	43
1977	2.9	2.6	9.6	−2.3	2.8	43
1978	2.5	2.9	8.8	−11.2	3.6	44
1979	2.2	3.0	8.6	−33.7	6.3	44
1980	1.2	1.6	8.3	−14.2	4.0	44
1981	0.4	0.7	6.4	−12.6	3.6	45
1982	−1.6	−0.6	6.4	−12.5	4.0	45
1983	−0.7	0.3	10.0	−13.4	4.4	45
1984	2.0	2.4	8.2	−4.2	2.6	45
1985	1.4	2.0	6.2	−9.4	3.1	45
1986	2.6	2.4	10.6	−5.8	3.3	45
1987	2.7	2.5	10.6	−4.0	3.0	45
1988	1.5	2.7	10.1	−16.5	5.2	45
1989	1.4	2.1	8.4	−15.3	4.5	45

Year	*Average*	*Mean*	*Max.*	*Min.*	*Standard deviation*	*Obs.*
1990	1.4	1.8	8.0	−7.6	3.5	45
1991	0.8	1.2	10.6	−12.1	4.1	47
1992	1.4	1.0	9.9	−5.4	3.4	48
1993	1.3	1.5	5.8	−2.8	2.4	49
1994	2.6	2.8	9.8	−6.3	2.7	49
1995	2.8	2.5	17.1	−7.9	3.5	49
1996	2.5	2.3	7.9	−2.3	2.2	51
1997	3.7	3.4	12.2	−0.6	2.2	51
1998	2.4	3.0	7.6	−6.6	2.6	51
1999	1.6	2.6	9.5	−8.1	3.7	51
2000	3.0	3.2	8.1	−4.4	2.4	51
2001	0.7	1.0	6.6	−7.3	2.4	51
2002	0.8	1.2	6.6	−12.5	3.5	51
2003	1.6	1.5	7.8	−9.9	2.6	51
2004	3.6	3.2	15.0	−0.2	2.4	51
2005	3.2	2.5	9.6	0.2	2.2	51
2006	3.9	3.4	10.5	1.4	2.1	51
2007	3.8	3.4	10.1	0.4	2.3	51
2008	1.1	1.1	7.7	−5.2	2.6	51
2009	−3.7	−3.9	3.8	−15.7	3.4	51
2010	2.5	1.9	10.6	−5.3	3.0	51
2011	2.3	2.1	8.6	−9.0	2.8	51
2012	0.6	0.8	8.1	−6.5	2.7	51
2013	1.0	1.0	11.6	−17.1	3.5	51
2014	1.5	1.4	5.9	−5.5	1.8	48
All	2.1	2.3	17.1	−33.7	3.6	2,496

Source: Own elaboration.

Annex 4. Derivation of the Relation between Prices, Money, Debt, and Inflation

As proposed by Kwon et al. (2009), a simplified version of Castro et al. (2003) can be used to derive a functional relation between the price level, money, debt and output. In said version, a representative consumer is endowed with fixed resources *(y)* for each period, and allocates their real wealth among real consumption *(c)*, real domestic money $\left(m/p\right)$, and non-indexed real government bonds $\left(b/p\right)$ in order to maximize the following utility function:

A.1 $$\sum_{t=0}^{\infty}\beta^{t}\left(\ln\left(c_t\right)+\gamma\ln\left(\frac{m_t}{p_t}\right)\right).$$

Subject to a resource constraint given by

A.2 $$c_t+\frac{m_t}{p_t}+\frac{b_t}{p_t}=y_t-\tau_t+\frac{m_{t-1}}{p_t}+\frac{i_{t-1}b_{t-1}}{p_t},$$

where τ is the fixed lump-sum tax and i_{t-1} is a nominal gross return of a government bond between periods $t-1$ and t. This maximization problem yields the following standard first-order conditions for consumption and real money demand, respectively:

A.3 $$\frac{c_{t+1}}{c_t}=\frac{\beta i_t}{\pi_{t+1}},$$

A.4 $$\frac{m_t}{p_t}=\frac{\gamma c_t i_t}{i_t-1},$$

where $\pi_t=p_{t+1}/p_t$. These two first order conditions nest a Cagan-type money demand function that is inversely related to inflation expectations.

The government is faced with the following intertemporal budget constraint:

$$\text{A.5} \quad G_t + (i_{t-1} - 1)\frac{B_{t-1}}{p_t} = \tau_t + \frac{(M_t - M_{t-1})}{p_t} + \frac{(B_t - B_{t-1})}{p_t}.$$

Forward iteration on Equation A.5 and no-Ponzi game conditions on the government imply the following long-term constraint of the government:

$$\text{A.6} \quad \frac{i_{t-1}B_{t-1}}{p_t} = \sum_{j=0}^{\infty}\frac{\tau_{t+j}}{R_{t,j}} - \sum_{j=0}^{\infty}\frac{G_{t+j}}{R_{t,j}} + \sum_{j=0}^{\infty}\frac{M_{t+j} - M_{t+j-1}}{p_{t+j}R_{t,j}},$$

where G is real government spending and $R_{t,j}$ is the compounded real discount rate expressed as

$$R_{t,j} = \prod_{h=1}^{j} r_{t+h}$$

where r_{t+h} is the exogenous real interest rate between periods $t+h-1$ and $t+h$. In the case of a fiscal policy rule where part of the debt service $(1-\delta)$ is covered with future primary surpluses and by monetizing the remainder (δ), we obtain the following money supply function:

$$\text{A.7} \quad \frac{M_t}{P_t} = \frac{i_t - 1}{i_t}\left[\frac{\delta i_{t-1}B_{t-1}}{p_t} + \frac{M_{t-1}}{p_t} - \sum_{j=1}^{\infty}\frac{M_{t+j}}{p_{t+j}R_{t,j}}\frac{i_{t+j} - 1}{i_{t+j}}\right].$$

Equation A.7 shows that the path of money supply is determined by the extent of debt monetization (the first variable in parenthesis on the right) and savings in the future interest payments brought about by current monetary financing of the budget deficit (third variable in parenthesis on the right).

Imposing equilibrium conditions on Equations A.4 and A.7, and exploiting the recursive nature of the Euler equation in A.3, we obtain the equilibrium price as follows:

A.8
$$p_t = \frac{(1-\beta)(M_{t-1} + \delta i_{t-1} B_{t-1})}{\gamma c_t}.$$

Given the recursive nature of the equilibrium and no arbitrage between bond and real asset returns $(r_{t+1} = i_t / \pi_t)$, the equilibrium price can be rearranged to:

A.9
$$p_t = \frac{(1-\beta)(M_t + \delta B_t)}{\gamma c_t}.$$

Using real income through real GDP (w) as a proxy variable for consumption in each period t, (c_t), results in Equation A.9 being equivalent to Equation 1.

Annex 5. Countries Analyzed and their Classification into Developed, Developing, Net Creditor and Net Debtor

Country	*Developing*		*Developed*
	Net debtor	*Net creditor*	*Net creditor*
Argentina	X		
Australia			X
Austria			X
Barbados	X		
Belgium			X
Bolivia		X	
Brazil	X		
Canada			X
Chile	X		
Colombia	X		
Costa Rica	X		
Czech Republic			X
Denmark			X
Dominican Republic	X		
Ecuador	X		
El Salvador	X		
Estonia			X
Finland			X
France			X
Germany			X
Greece			X
Guatemala	X		
Honduras	X		
Hungary	X		
Iceland			X

Ireland			X
Israel			X
Italy			X
Jamaica	X		
Japan			X
Korea			X
Luxemburg			X
Mexico	X		
Netherlands			X
New Zealand			X
Nicaragua	X		
Norway			X
Panama	X		
Paraguay	X		
Peru	X		
Poland	X		
Portugal			X
Slovakia			X
Slovenia			X
Spain			X
Sweden			X
Switzerland			X
Turkey	X		
United Kingdom			X
United States			X
Uruguay	X		
Venezuela		X	

Source: Own elaboration based on the World Economic Outlook (WEO) 2014.

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Anchoring of Inflation Expectations in Mexico

Marco A. Acosta

Abstract

This study analyzes short, medium and long run inflation expectations anchorage among professional forecasters from the private sector in Mexico before and after the financial crisis of 2008 by introducing a novel classification that catalogs to a large extent the econometric efforts that have been made for its measurement. The three dimensions covered by this classification are sensitivity, resilience and credibility. The results show that for the period evaluated after the 2008 financial crisis and as the horizon for which inflation forecasts are made increases, expectations are better anchored.

Keywords: inflation expectations, anchorage, sensitivity, resilience, credibility.

JEL classification: C12, C13, E31, D84.

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1. INTRODUCTION

The monetary policy of Banco de México aims to influence interest rates in order to bring price behavior into line with the path of inflation towards its long run target. Inflation expectations are therefore of utmost importance given that forecasts regarding the future costs and income of economic agents are crucial for setting the prices of the goods and services they supply. The greater the public's trust in the central bank, the better expectations will be anchored, which translates into an environment of low and stable inflation that in turn fosters conditions favoring sustained economic growth.

This paper analyzes the anchorage of inflation expectations among professional forecasters from the private sector at different horizons from January 2002 to May 2017, and for two subperiods divided by the 2008 financial crisis, using linear regressions and vector autoregressive (VAR) models. In specific, I assess three dimensions with respect to the anchoring of inflation expectations: *1)* the sensitivity of medium and long-term expectations to contemporary inflation and short-term expectations; *2)* resilience to inflation shocks; and *3)* the credibility of Banco de México. Documents found in the literature usually only focus on one of these three dimensions, naming the dimension they assess anchoring. This study therefore integrates the literature and categorizes existing types of anchoring in order to provide coherence to the findings.

The outcomes show how the behavior of inflation expectations has been consistent with the process of convergence towards low and stable inflation during recent years. It shows how the distribution of expectations has been centered around the permanent 3% target for inflation and the upper limit of the variability interval. Moreover, dispersion is modest, and bias is not statistically different from zero in most of the periods.

As for sensitivity, the paper shows that short-term inflation expectations, defined as those for the following 12 months, are associated to changes in the contemporaneous inflation process. Medium-term expectations, which encompass forecasts from one- to four-year ahead, are less affected than short-term ones; while long-term expectations forecasting five- to eight-year ahead do not experience any effects. It also shows that long-term inflation expectations are not affected by short-term ones.

With respect to resilience, the outcomes reveal that inflation shocks do not influence the formation of expectations under the current economic setting, even including expectations over shorter horizons such as those for 12 months. It can also be seen that resilience coefficients for the estimations are not statistically significant for the periods before or after the 2008 financial crisis, revealing the stability of the inflation process since the start of the last decade.

The evidence suggests that credibility in the central bank's long-run inflation target grows as the horizon for which inflation forecasts are made increases. The credibility of implicit inflation derived from an autoregressive vector exercise displays a similar behavior: the longer the forecast horizon, the more credible the inflation target becomes. The aforementioned could be due to the capacity the central bank has demonstrated to respond to inflation shocks with the monetary policy and communication tools at its disposal in order to bring inflation into line with the 3% target.

The exercises for the periods before and after the 2008 financial crisis show that inflation expectations are better anchored at all forecast horizons for the dimensions of sensitivity and credibility postcrisis. As for the resilience indicator, expectations do not appear to have been affected in the pre- or postcrisis periods.

The rest of the paper is organized as follows. Section 2 presents the development of achievements in inflationary matters from 1994 to date. Section 3 describes the dimensions in which the anchoring of inflation expectations is analyzed. Section 4 presents an analysis of the data employed, particularly examining the dispersion, skewness and rationality of expectations.

The Section 5 describes the outcomes. Finally, concluding remarks are provided.

2. TRANSITION TOWARDS LOW AND STABLE INFLATION IN MEXICO

On account of the 1994-1995 crisis, Mexico adopted a set of measures aimed at maintaining inflation at low and stable levels. Among these stands out the establishment of a target for the current accounts commercial banks hold at the central bank, commonly known as the *short*, a tool that allows for controlling liquidity in the economy with the aim of eliminating inflationary pressures. In 1998, Banco de México accompanied its announcements of changes in the short with a discussion of the main reasons for such modifications, thereby making the application of monetary policy transparent. Subsequently, in the year 2000, the Bank began publishing quarterly inflation reports and in 2001 the process towards transparency was boosted by announcement of the adoption of an inflation targeting regime.[1,2]

The successful reduction of inflation in Mexico has been well documented due to the short time it took. Triple and double-digit inflation had been recorded in the eighties and nineties respectively, but after 2000 it fell to just single digits. Furthermore, as described by Chiquiar et al. (2007), inflation acquired important statistical properties: in specific, it switched from a nonstationary to a stationary process around the end of 2000 and the beginning of 2001. From an economics point of view, statistical behavior implies that shocks to inflation become diluted over time and do not generate second round effects that could alter the price formation process of

1 For an in-depth discussion on the transition towards an inflation targeting regime see Ramos-Francia y Torres (2005).

2 The works of Bernanke *et al.* (1999) and Corbo *et al.* (2001) illustrate the favorable behavior of inflation in countries with an inflation targeting regime as compared to other regimes.

the economy. Moreover, Acosta (2018) employs a quantile regression with structural changes approach to show that after the year 2000 inflation follows a stationary behavior in all its conditional quantiles.

Another important change is that inflation in Mexico became a mostly time-dependent process, which allows revisions to be made that do not depend on the state of the economy, allowing for better planning among the agents involved (see Gagnon, 2009). A downward flexibility in prices has also been observed during recent years as shown by Cortés et al. (2011) on the basis of the microdata used for calculating the national consumer price index. The majority of price revisions had previously been upwards.

Inflation's interaction with other macroeconomic variables that can influence it has also changed. Capistrán et al. (2011) and Cortés (2013) found a reduction in the pass-through of exchange rate fluctuations to inflation in the period after the inflation target was adopted. The aforementioned might respond to the absence of any second-round effects from international commodity price variations and the lack of any permanent effects on inflation from tax changes such as those implemented in 2010, as mentioned by Aguilar et al. (2014).

With respect to inflation expectations, the topic studied in this paper, the work on Mexico by García-Verdú (2012) stands out. The latter employs the model of Mankiw et al. (2003) to explore the dispersion of inflation expectations among professional forecasters from the private sector. The model of Mankiw et al. (2003) is based on the principle that there are costs implicit in collecting and processing information for readjusting inflation forecasts, meaning only some economic agents update them. This leads to dispersion between the expectations of agents who use recent and lagged data. The findings of García-Verdú (2012) show that a larger proportion of forecasters from the private sector update their inflation expectations, which coincides with lower levels of dispersion observed in the data. Likewise, García-Verdú (2012) study the dispersion and skewness of expectations and determine that they have diminished, which they attribute to a more stable environment and the reduction of potential risks, respectively.

3. DIMENSIONS FOR THE ASSESSMENT OF THE ANCHORING OF INFLATION EXPECTATIONS

If inflation expectations were perfectly anchored there would be no relation at all between actual inflation and economic agents' forecasts. Nevertheless, this level of anchorage is not usually seen in the data, but it allows for carrying out a test in which anchorage is defined by the level of linear dependence displayed by inflation expectations with respect to observed and lagged inflation. Among the papers that have characterized the anchorage of expectations in this way are those of Levin et al. (2004) and Ehrmann (2015). The same principle is applicable to medium and long-term expectations with respect to short-term ones; that is, if inflation expectations at more distant horizons are well-anchored they should be insensitive to changes in expectations at shorter horizons. This hypothesis accepts movements in short-term expectations, meaning they are not perfectly anchored. It also sets forth a scenario where medium and long-term expectations can be anchored if they do not respond to their short-term counterparts. In particular, Łyziak and Paloviita (2017) study said anchorage for the European Union. Anchoring tests with the previously mentioned characteristics will be referred to as sensitivity tests.

If inflation expectations are well-anchored, shocks to inflation should not affect them, given that economic agents expect the central bank to act in line with its long-run inflation target. Among the papers that have characterized the anchoring of inflation expectations with respect to the linear impact of an inflation shock are Mariscal et al. (2014) and Aguilar et al. (2014). Those studies employ a variable that takes a maximum value of between one and the difference between lagged inflation and its long-run target to define shocks to inflation. This type of tests shall be called resilience tests.

The anchoring of expectations for a central bank can be evaluated as the extent to which professionals from the private sector believe in the long-run inflation target. Bomfim and Rudebusch (2000) use a linear regression as a reference where the weighted sum of the long-term target and lagged values of inflation are made equal to inflation expectations in order to test said hypothesis. The coefficient given to the target is therefore the weight or degree of credibility professionals have in their central bank. Meanwhile, Demertzis et al. (2009) calculate the implicit anchoring of inflation expectations estimating a VAR model, using that methodology to assess whether implicit anchoring coincides with the long-run target for inflation. These measures are referred to as creditability.

4. DATA

Data employed in this paper is taken from Banco de México's Encuestas de los Especialistas en Economía del Sector Privado (EEBM, Surveys of Forecasters on Economics from the Private Sector), which has been conducted on a monthly basis since September 1994 and includes forward-looking questions on economic matters aimed at obtaining expectations regarding important macroeconomic variables such as the exchange rate, interest rates, wages and inflation, among others.[3] The collected information is used to prepare a monthly report that is published at the start of each month and shows the consensus of professionals' forecasts for each variable and time horizon. Said consensus is represented by the average and median of the forecasts.

[3] In the period studied 86, 68 and 59 institutions or individuals participated answering questions on their short, medium and long-term inflation expectations, providing an average of 30, 28 and 27 answers to each survey, respectively.

This paper analyzes the medians of inflation expectations at three time horizons because they better capture the consensus of economic forecasters as an extreme value could substantially alter the average, without changing that of the median.[4] The short-term horizon refers to the forecasts professionals make 12-month ahead for annual inflation; the medium-term includes forecasts made four- year ahead; while the long-term considers the forecasts of economic agents for a time interval of five- to eight-year ahead.

Figure 1 shows the performance of headline inflation observed during the study period and expectations for it at the three time horizons specified above. The series have different starting points because the EEBM began to ask questions regarding medium and long-term expectations in January 2004 and August 2008, respectively. Although for short-term expectations the EEBM contains data available for periods before January 2002, I decided to begin on that date because it is the first full year in which inflation follows a stationary path.[5]

Although the medians of answers are taken as the consensus among professionals, it is important to test whether the median actually does represent the central tendency of the answers and whether they are converging towards the target.[6] To that end, I analyze empirical density functions, dispersion and skewness of forecast data, as well as its rationality. With respect to density, Figure 2 presents the empirical distributions of inflation expectations at different horizons. Expectations for 12-month ahead are mostly concentrated in the 3% to 6% interval. Nonetheless, it can be seen how densities shifted to the left, towards the long-run inflation target, as time progressed, and in recent years it

[4] The anchoring of inflation forecasts at time horizons that may change, such as in the case of year-end inflation forecasts, are not studied.

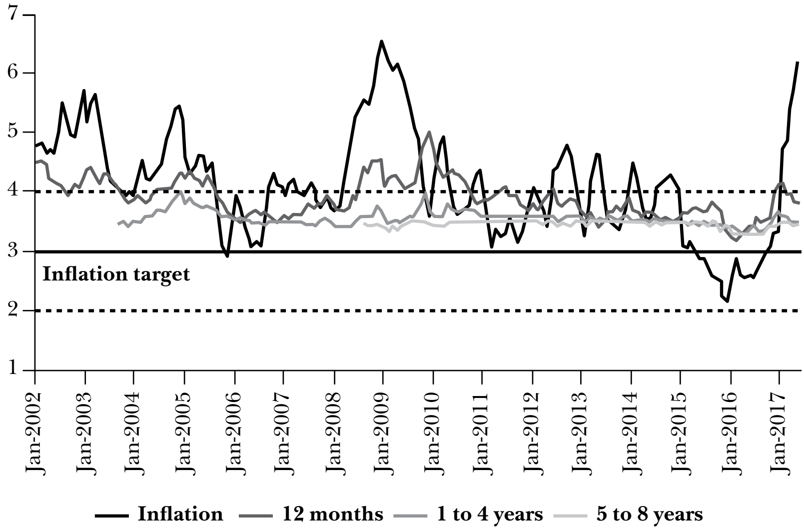
[5] Chiquiar *et al.* (2007) point out that in December 2000 and April 2001, headline as well as core inflation underwent a structural change shifting from a nonstationary to a stationary process.

[6] Carrera (2012) uses histograms to show that inflation expectations in Peru are centered.

Figure 1

SHORT-, MEDIUM- AND LONG-TERM EXPECTATIONS

Annual variation in percentages

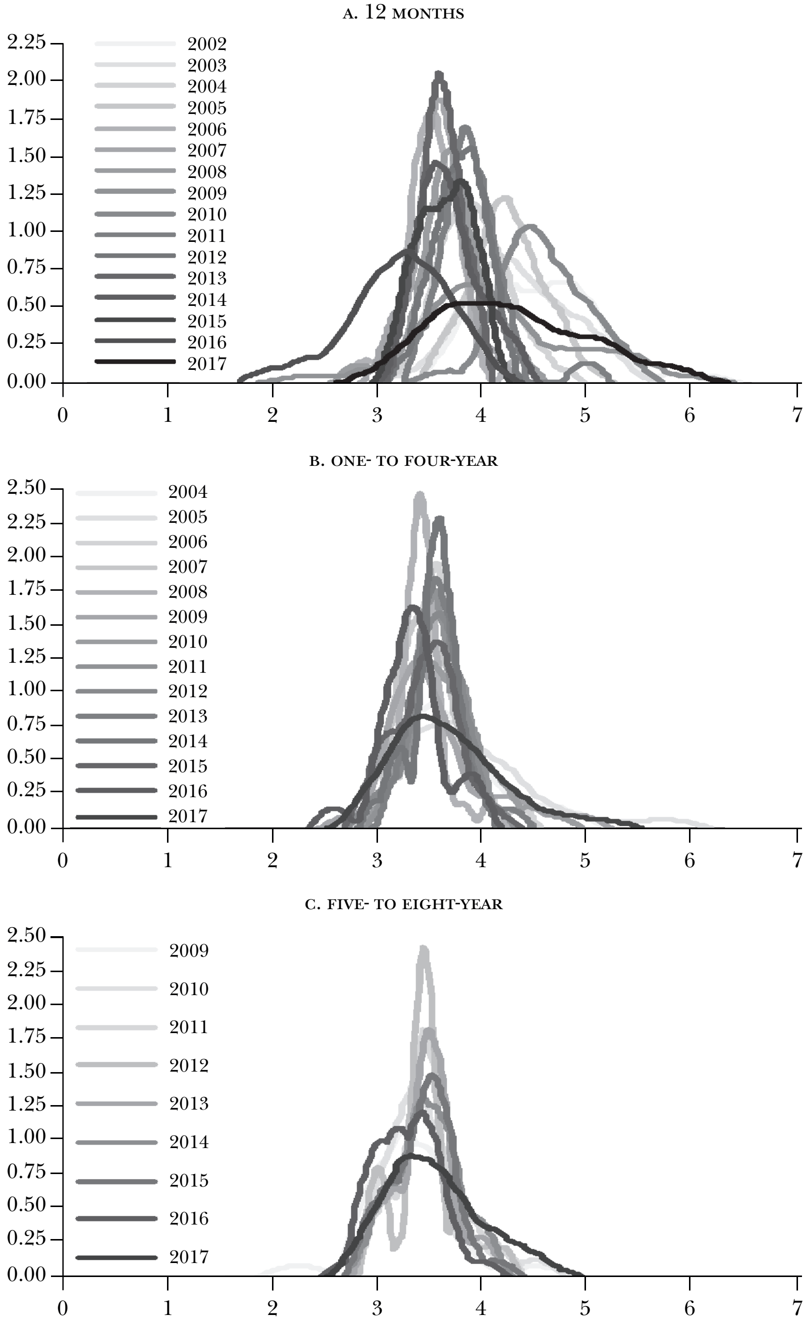


is located in a narrower interval, between 3% and 4.5%. Inflation expectations for one- to four-year ahead are concentrated between 3% and 4.5%, while long-term ones are centered between the 3% inflation target and the upper bound of the variability interval.

Dispersion, calculated as the month to month interquartile range inside which economic agents specified their expectations, is low (Figure 3). Said characteristic is key for assessing anchorage given that a smaller dispersion implies greater agreement among professionals. In particular, on average, the interquartile ranges of inflation expectations from shorter to longer horizons are 54, 34 and 34 basis points. Moreover, it can be seen that during periods of high economic uncertainty dispersion increases at all horizons, a characteristic clearly observable between 2008 and 2010 (Figure 3d). Nevertheless, this growth is modest and temporary, evidence of rigidity among professionals to change their forecasts.

Figure 2

EMPIRICAL DISTRIBUTION OF THE SPECIALISTS' EXPECTATIONS[1]

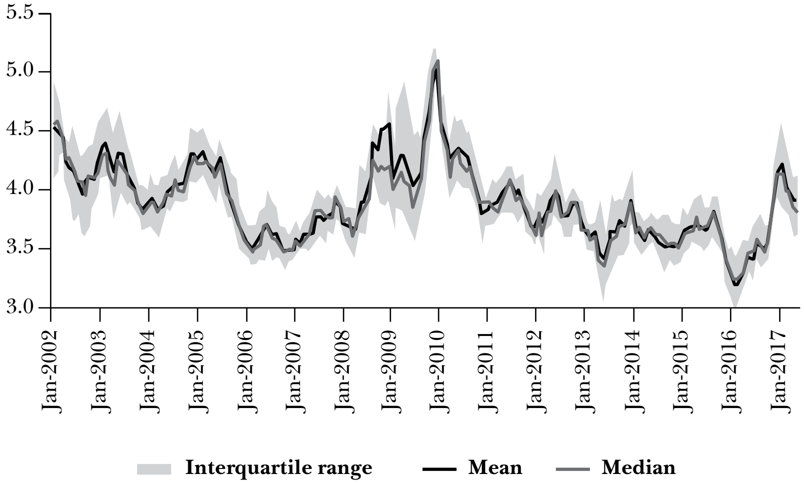


[1] Expectations correspond to January of each year.

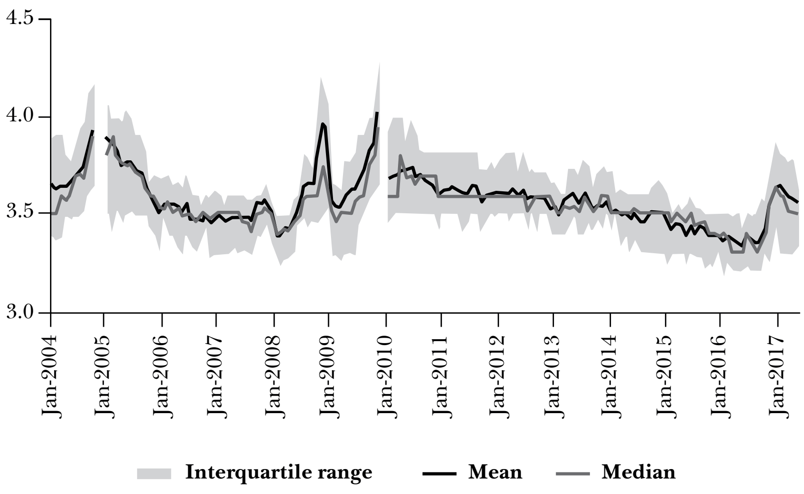
Figure 3

DISPERSION, MEAN, AND MEDIAN OF THE SPECIALISTS' EXPECTATIONS[1]

A. 12-MONTH EXPECTATIONS



B. ONE- TO FOUR-YEAR EXPECTATIONS

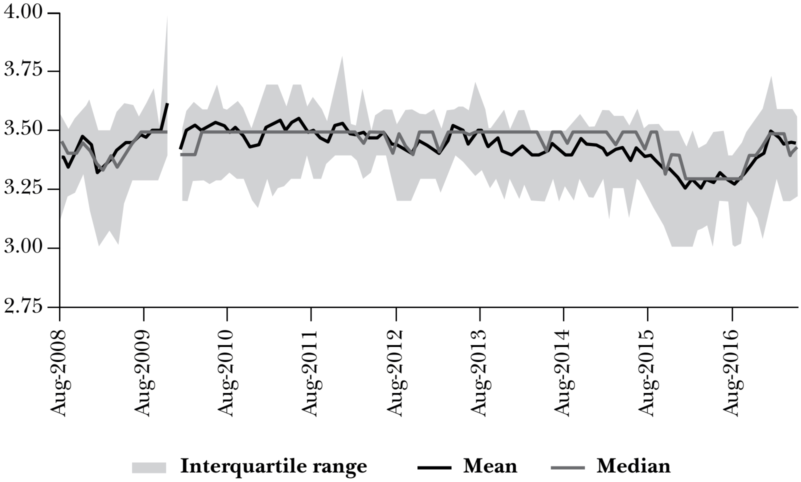


[1] Expectations correspond to January of each year.

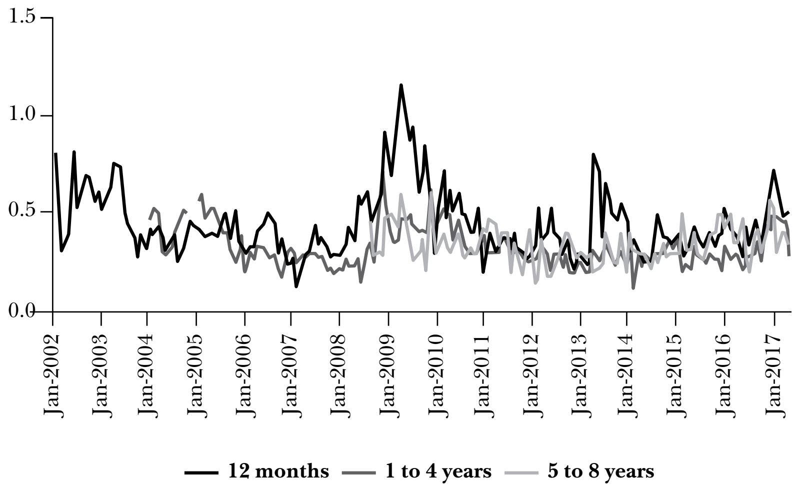
Figure 3 (cont.)

DISPERSION, MEAN, AND MEDIAN OF THE SPECIALISTS' EXPECTATIONS[1]

c. 8-year expectations



d. dispersion to different horizons



[1] Expectations correspond to January of each year.

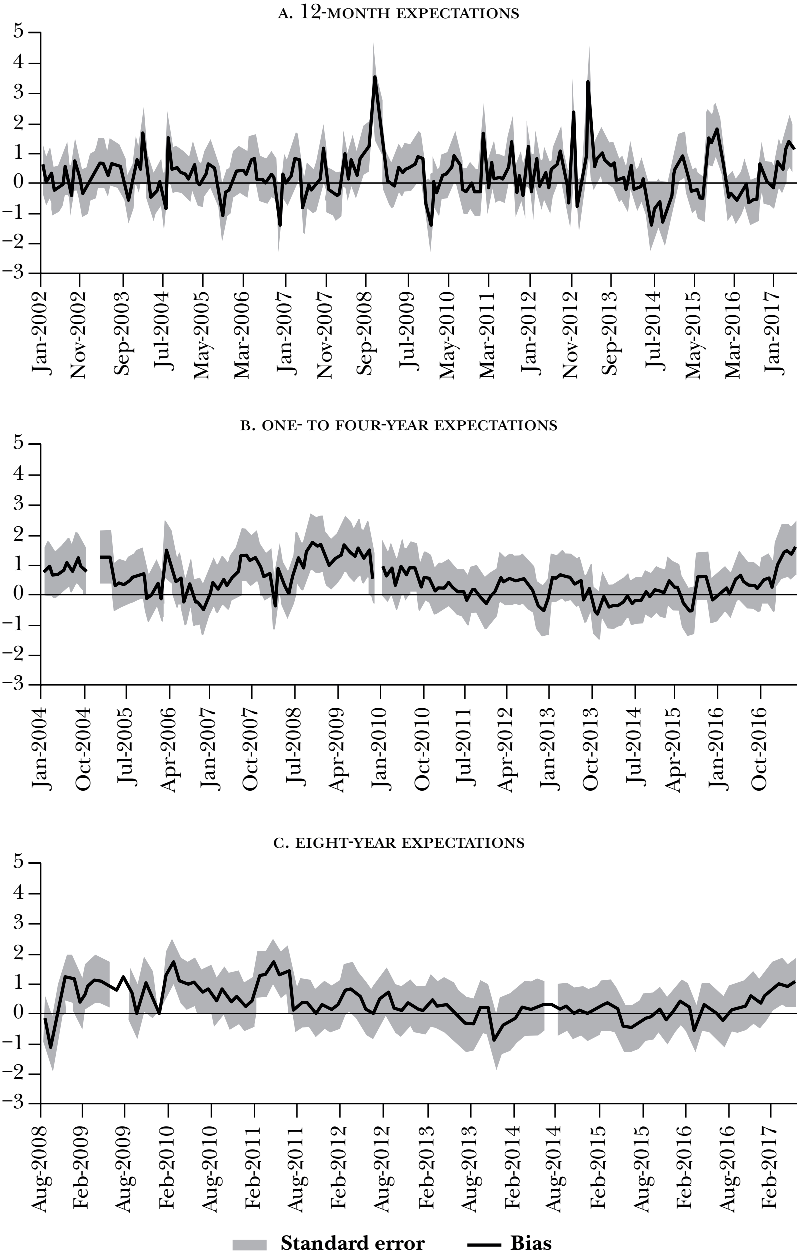
Bias is interpreted as the existence of upward risks if its value is positive and downward risks if it is negative. Expectations at all horizons appear to exhibit neutral risk; that is, their bias is not statistically significant for the majority of periods (Figure 4). Nevertheless, for medium and long-term horizons there appear to be consecutive data sets in which professionals forecast upward risks characterized by positive biases (Figures 4b and 4c) that coincide with periods of greater volatility. Hence, it is possible to see periods where inflation expectations experienced higher uncertainty represented by upward risks in the inflation process. Nonetheless, this was not the case for the majority of periods which presented null skewness and low levels of volatility.

Many papers focus on exploring the coherence between inflation expectations and the rational expectations hypothesis, understood as the impossibility of obtaining predictable errors in the forecasts. To explore whether inflation expectations fulfill the defined characteristic, I perform a set of tests commonly used in the literature and reported in Mankiw et al. (2003) for the case of the United States.

Table 1 presents the results of the tests of expectation rationality. Panel A reports these results, regressing forecast errors on a constant. This is a simple test to evaluate whether inflation expectations are centered on the correct value. The value of the constant is not significant, meaning the forecast errors of professionals are therefore centered on the correct value. Panel B tests whether there is information available in these expectations that can be used to predict forecasting errors. The null hypothesis is that the regression should have no predictive power. As can be seen, the null hypothesis is rejected, meaning there is information that can be exploited. Panel C tests whether today's errors can be forecasted based on yesterday's errors; that is, if there is autocorrelation. The coefficient associated with autocorrelation is not statistically significant. Finally, Panel D assesses whether inflation expectations take account of available macroeconomic information to make the forecasts. The null hypothesis is that macroeconomic variables should not help to predict forecasting errors. However, the null hypothesis is rejected because all the macroeconomic variables help to improve the forecasts.

Figure 4

BIAS OF INFLATION EXPECTATIONS



In sum, the medians are a good indicator for the central tendency of inflation expectations from the private sector. Dispersion is modest, and in most periods, skewness is not statistically different from zero. As for the rationality of expectations, the forecasts are not efficient because they do not leverage all the information from previous periods or available macroeconomic data. Nevertheless, they do not exhibit bias and forecast errors diminish over time. For this reason, median inflation expectations are the indicator recommended as a measure of the central tendency of the data for performing an assessment of the anchoring of inflation expectations.

Table 1

TEST OF FORECAST RATIONALITY

A. Skewness test $\pi_t - \pi^e_{t\|t-12} = \alpha$					
α	0.05 (0.15)				
B. Is information in the forecast fully exploited? $\pi_t - \pi^e_{t\|t-12} = \alpha + \beta\pi^e_{t\|t-12}$					
α	2.94[a] (0.54)	β	−0.71[a] (0.11)		
$H_o: \alpha = \beta = 0$		value p= 0.00			
C. Are forecasting errors persistent? $\pi_t - \pi^e_{t\|t-12} = \alpha + \beta\left(\pi_{t-12} - \pi^e_{t\|t-24}\right)$					
α	0.07 (0.46)	β	0.10 (0.15)		
D. Are macroeconomic data fully exploited? $\pi_t - \pi^e_{t\|t-12} = \alpha + \beta\pi^e_{t\|t-12} + \gamma\pi_{t-13} + \kappa CETES + \xi IGAE$					
α	4.92[a] (0.80)	β	−1.09[a] (0.24)	γ	−0.42[a] (0.11)
κ	0.24[a] (0.04)	ξ	−0.13[a] (0.03)		
$H_o: \gamma = \kappa = \xi = 0$		value p= 0.00			

Note: [a], [b] and [c] denote statistical significance at the 10%, 5% and 1% levels, respectively.

5. EMPIRICAL ANALYSIS

In a monetary policy credibility framework, deviations of inflation from its long-term target should be transitory. Thus, economic agents should perceive observed deviations as something transitory that will converge to its target over the long-run and remain there. Nevertheless, there are different risks due to which economic agents' expectations regarding inflation might undergo changes that include: contamination of medium and long-term expectations due to modifications in contemporaneous inflation or short-term expectations (sensitivity), inflation shocks negatively influencing the behavior of expectations at all horizons (resilience) or a central bank that is more tolerant of deviations from its long-run target (credibility). It is therefore important to monitor inflation expectations to enable early detection of any adverse effects in them. Thus, this empirical analysis presents a complete methodology for evaluating expectations in order to identify and classify the type of impact expectations could undergo.

5.1 Sensitivity of Expectations

The sensitivity of inflation expectations is assessed in two different ways in this paper. The first consists of assessing whether changes in the contemporaneous inflation process impact inflation expectations in line with Ehrmann (2015). Thus, short-term expectations are expected to be strongly affected, medium-term ones affected to a lesser extent than short-term ones, while long-term expectations are not affected at all. The second evaluation highlights that medium and long-term inflation expectations should not be affected by changes in short-term ones. The methodology employed follows that specified by Łyziak and Paloviita (2017).

5.1.1 Relation of Inflation Expectations with Respect to Lagged Inflation

If medium and long-term expectations are well-anchored they should not be affected at all by movements in lagged inflation, while short-term ones can be affected by the lagged inflation process. To test this assertion following the methodology of Ehrmann (2015), I estimate

$$\pi^e_{t|t+n} = \alpha + \beta\pi_{t-1} + \varepsilon_t, \tag{1}$$

where $\pi^e_{t|t+n}$ is inflation expectations formed in period t at the forecast horizon $t+n$; π_{t-1} is lagged inflation; α is the regression constant; β is the lagged inflation coefficient; and ε_t is the regression error. If β is not significant or very close to zero it would indicate that expectations are not contaminated by the inflation process.

Given that the anchoring of expectations might have undergone changes due to the reduced global demand stemming from the 2008 financial crisis I estimate

$$\pi^e_{t|t+n} = (1-CF)(\alpha_{ACF} + \beta_{ACF}\pi_{t-1}) + CF(\alpha_{DCF} + \beta_{DCF}\pi_{t-1}) + \varepsilon_t. \tag{2}$$

The variable CF represents the 2008 financial crisis and takes a value of zero for each of the periods before April 2008 and one for subsequent periods just as in Łyziak and Paloviita (2017). To test for robustness, equation 1 is estimated with six-year rolling windows.

Table 2 shows the outcomes of equations 1 and 2. The lagged inflation coefficient (β) for the full sample of 12-month ahead expectations is significant and takes the value of 0.22, which leads to adjustments in expectations after changes in observed inflation. Meanwhile, for medium and long-term expectations said coefficient is small and only significant for four-year ahead expectations; that is, actual inflation does not appear to affect expectations at longer horizons.

Coefficient β for expectations in periods after the 2008 financial crisis exhibits a substantial reduction. In particular, the coefficient for 12-month ahead expectations shift from 0.31 to 0.19, and for one to four-year ahead expectations it decreases from 0.18 to 0.04, the spread being statistically significant in both cases (Table 2).

Table 2

RELATION OF INFLATION EXPECTATIONS WITH RESPECT TO LAGGED INFLATION

	β	R^2	β_{ACF}	β_{DCF}	R^2	$H_0 : \beta_{ACF} = \beta_{DCF}$
Twelve-month ahead expectations	0.22[a]	0.36	0.31[a]	0.19[a]	0.38	2.71
Four-year ahead expectations	0.06[a]	0.20	0.18[a]	0.04[b]	0.35	2.93
Eight-year ahead expectations	0.01	0.03	na	na	na	na

Note: Ordinary least square estimates were performed with Newey-West standard errors. [a] and [b] denote statistical significance at the 1% and 5% level, respectively. The value reported for the hypothesis test is the t statistic. na stands for not available.

Figure 5a shows the lagged inflation coefficient of the six-year rolling window regressions, which diminished from May 2008 to May 2017, reaching statistically nonsignificant values after June 2015. Meanwhile, Figures 5b and 5c illustrate that although the lagged inflation coefficient for medium and long-term expectations increased between 2015 and 2016, it exhibited relatively small values. The aforementioned is consistent with that seen in Łyziak and Paloviita (2017) for periods after the 2008 financial crisis.

The increase in the sensitivity of medium and long-term expectations seen in the later periods could be explained by the volatility of energy prices in Mexico stemming from a regime change they have undergone since the energy reform. In specific, the initial falls in energy prices observed at the start of 2015 appear to have pushed long and medium-term expectations downwards. These moved closer to the long-run inflation target at the end of 2015 when headline inflation was below target. Another possible explanation is the increase in exchange rate volatility caused by the start of electoral campaigning in the United States (USA). In particular, from June 2015 to November 2016 (the start of campaigning up until when the elections are held in the USA), the Mexican peso depreciated around 25%. Nevertheless, the increased sensitivity observed in medium and long-term expectations appears to have been temporary, with even a slight downward trend being seen in the coefficient associated to sensitivity during the later periods (Figures 5b and 5c).

5.1.2 Sensitivity of Medium and Long-term Inflation Expectations to Short-term Ones

If inflation expectations are well-anchored, medium and long-term expectations should not respond to movements in short-term ones. To examine said relation I use the methodology proposed by Łyziak and Paloviita (2017). In particular, I estimate

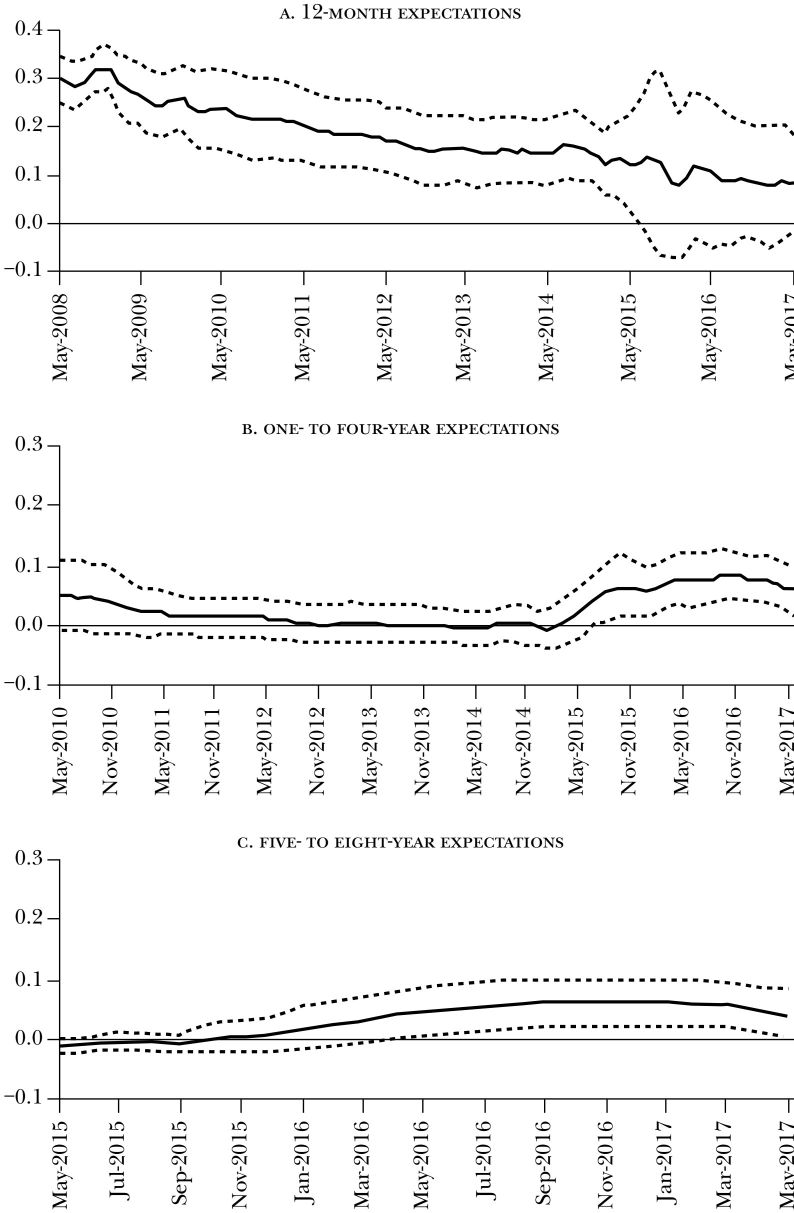
$$\pi^e_{t|t+n} = \alpha + \lambda \pi^e_{t|t+m} + \varepsilon_t, \quad (3)$$

where $\pi^e_{t|t+n}$ refers to inflation expectations formed in period t for the forecast horizon $t+n$; $\pi^e_{t|t+m}$ is inflation expectations formed in period t for the forecast horizon $t+m$; α is the regression constant; λ is the lagged inflation coefficient; and ε_t is the regression error.

Figure 5

RELATION OF INFLATION EXPECTATIONS TO LAGGED INFLATION

Six-year rolling windows (coefficient β)



It is important to mention that $t+n>t+m$, given that the dependent variable are medium and long-term expectations. If the coefficient λ is not significant or close to zero it indicates that long-term inflation expectations are insensitive to fluctuations in short-term ones. Due to the fact that the 2008 crisis could have affected the relation between expectations I estimate

$$\pi^e_{t|t+n} = (1-CF)\left(\alpha_{ACF} + \lambda_{ACF}\pi^e_{t|t+m}\right) + CF\left(\alpha_{DCF} + \lambda_{DCF}\pi^e_{t|t+m}\right) + \varepsilon_t. \quad (4)$$

With these equations it is possible to estimate how long-term expectations respond to adjustments in short-term ones. To identify any possible changes in the coefficient of short-term expectations I estimate equation 3 with six-year rolling windows.

Expectations for one- to four-year ahead exhibit a significant, although relatively small coefficient, which translates into a modest impact deriving from the behavior of short-term expectations. Furthermore, the coefficient decreases after the 2008 financial crisis, to be specific, it shifted from 0.49 to 0.23 (Table 3). Meanwhile, long-term expectations do not respond to movements in short-term ones, which can be interpreted as a better anchoring of inflation expectations (Table 3).

Table 3

RELATION OF LONG-TERM INFLATION EXPECTATIONS WITH SHORT-TERM ONES

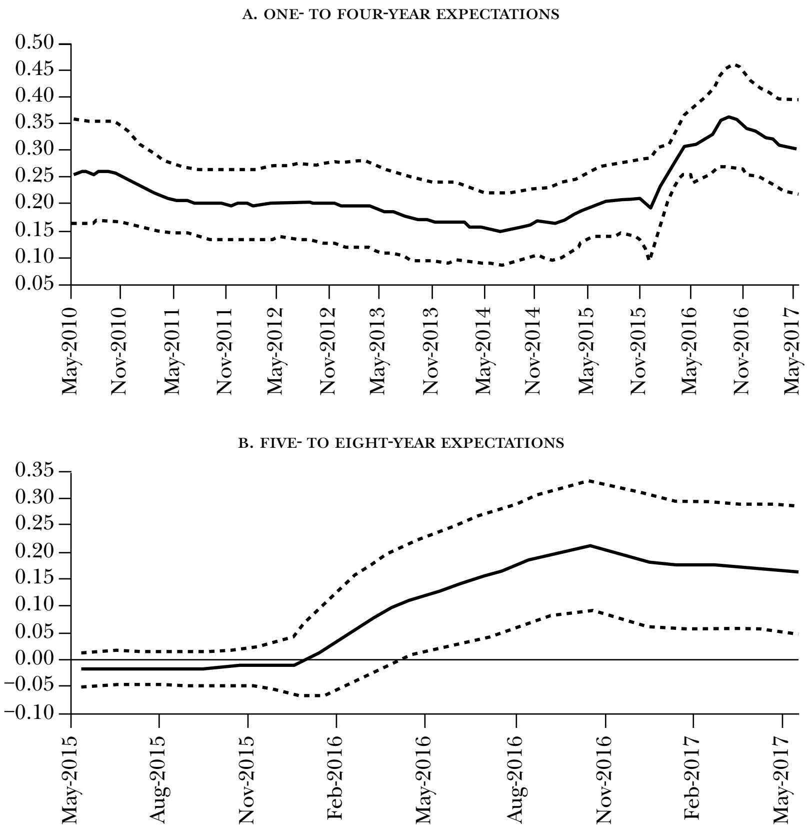
	λ	R^2	λ_{ACF}	λ_{DCF}	R^2	$H_0: \lambda_{ACF} = \lambda_{DCF}$
Four-year ahead expectations	0.28[a]	0.52	0.49[a]	0.23[a]	0.64	3.32
Eight-year ahead expectations	0.05	0.06	na	na	na	na

Note: Ordinary least square estimates were performed with Newey-West standard errors. [a] and [b] denote statistical significance at the 1% and 5% level, respectively. The value reported for the hypothesis test is the *t* statistic. na stands for not available.

Figure 6

RELATION OF LONG-TERM TO SHORT-TERM INFLATION EXPECTATIONS

Six-year rolling windows (coefficient λ)



Figures 6a and 6b depict the coefficient λ associated to six-year rolling window regressions. There is a rebound in both expectations in December 2015, while in medium-term ones the coefficient increases, in long-term ones it shifts from being statistically nonsignificant to significant. Short-term expectations are affected by current inflation, meaning the recent instability of energy prices and exchange rate volatility have probably caused a similar effect to that described in the previous section for medium and long-term expectations.

5.2 Resilience Expectations to Inflation Shocks

The effects of inflation shocks on expectations is captured as the impact caused by an increase that exceeds the upper limit of the long-term inflation target. Based on the methodology of Mariscal et al. (2014) for measuring the anchorage of inflation expectations and employed by Aguilar et al. (2014) to calculate the effect of inflation shocks, equation 1 can be modified by adding some variables and being written as:

$$\pi^e_{t|t+n} = \alpha + \beta\pi_t + \gamma\pi^e_{t-1|t+n} + \delta \max[\pi_{t-1} - \pi^{Obj}, 1] + \varepsilon_t. \tag{5}$$

In order to measure the impact of shocks on expectations. It is important to point out that lagged expectations are added to equation 1 to denote that the model focuses on the fluctuations of inflation expectations. The aforementioned can be more easily seen by rearranging equation 5 as

$$\pi^e_{t|t+n} - \gamma\pi^e_{t-1|t+n} = \alpha + \beta\pi_t + \delta \max[\pi_{t-1} - \pi^{Obj}, 1] + \varepsilon_t.$$

I also include the variable $\max\left[\pi_{t-1} - \pi^{Obj}, 1\right]$ that takes the value of lagged inflation minus the long-term target when said value is greater than one or one if not. In this way the added variable captures variations in periods when inflation exceeded the upper limit of the variability interval set for the long-run inflation target. Hence, δ is the coefficient associated to inflation shocks. To calculate whether there were more pronounced effects before or after the 2008 financial crisis I estimate

$$\begin{aligned} \pi^e_{t|t+n} = (1-CF)\Big(&\alpha_{ACF} + \beta_{ACF}\pi_t + \gamma_{ACF}\pi^e_{t-1|t+n} + \\ &+\delta_{ACF}\max\left[\pi_{t-1} - \pi^{Obj}, 1\right] + \varepsilon_t\Big) + (CF)\big(\alpha_{DCF} + \beta_{DCF}\pi_t + \\ &+\gamma_{DCF}\pi^e_{t-1|t+n} + \delta_{DCF}\max\left[\pi_{t-1} - \pi^{Obj}, 1\right] + \varepsilon_t\big). \end{aligned} \tag{6}$$

Table 4 shows the coefficient for the impact of inflation shocks on expectations. The coefficients are not statistically significant at all expectation horizons, except for 12-month ahead inflation expectations prior to the crisis. It is therefore possible to infer that inflation does not influence the formation of expectations under the current economic environment, even for expectations at shorter horizons such as those for 12 months ahead. Moreover, it is possible to observe that the resilience coefficients (δ) for the estimations are not statistically different for pre- or postcrisis periods, revealing the stability of inflation after it became a stationary process.

Table 4

RESPONSE OF EXPECTATIONS TO INFLATION SHOCKS

	δ	R^2	δ_{ACF}	λ_{DCF}	R^2	$H_0:\delta_{ACF}=\delta_{DCF}$
12-month ahead expectations	0.05	0.87	0.10[b]	0.04	0.88	0.94
Four-year ahead expectations	0.00	0.78	0.09	0.00	0.80	1.68
Eight-year ahead expectations	−0.02	0.60	na	na	na	na

Note: Ordinary least square estimates were performed with Newey-West standard errors. [a] and [b] denote statistical significance at the 1% and 5% level, respectively. The value reported for the hypothesis test is the *t* statistic. na stands for not available.

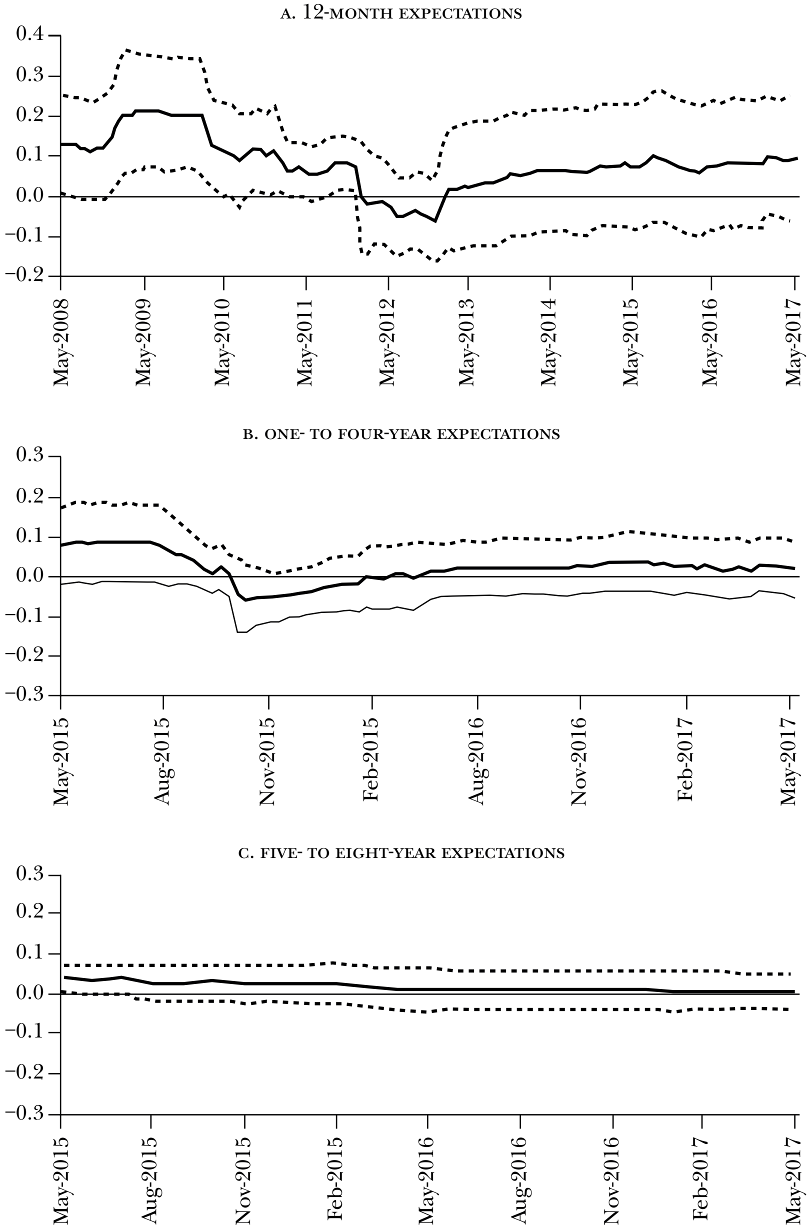
Using six-year rolling window regressions Figure 7a shows that during 2009 and up until the middle of 2010 short-term expectations were pushed upwards by fluctuations in inflation above the upper bound for the long-run inflation target. As of 2010 expectations remained insensitive to inflation shocks.

Figures 7b and 7c illustrate that medium and long-term expectations do not react to the spread between actual inflation and the upper limit of the inflation target given that during

Figure 7

RELATION OF EXPECTATIONS TO SHOCK ON INFLATION

Six-year rolling windows (coefficient δ)



most of the period the coefficient δ is not statistically significant, thus demonstrating that medium and long-term expectations are well-anchored and that inflation shocks do not affect them.

5.3 Credibility in Inflation Expectations

This paper measures the credibility of inflation expectations as the weight agents place on the central bank's long-run inflation target following the methodology of Bomfim and Rudebusch (2000). The analysis of credibility is also complemented by the VAR model proposed by Demertzis et al. (2008, 2009) which is used to calculate the anchorage and implicit credibility of inflation.

5.3.1 Credibility of Expectations with the Long-Run Inflation Target

This subsection examines how expectations are affected by the long-run target for inflation. The analysis uses the definition of Bomfim and Rudebusch (2000) for central bank credibility. In particular, the following equations are estimated:

$$\pi^{e}_{t|t+n} = \delta^{Obj}\pi^{Obj} + \left(1-\delta^{Obj}\right)\pi_{t-1} + \varepsilon_t, \tag{7}$$

$$\begin{aligned}\pi^{e}_{t|t+n} = {} & \left(1-CF\right)\left(\delta^{Obj}_{ACF}\pi^{Obj} + \left(1-\delta^{Obj}_{ACF}\right)\pi_{t-1}\right) + \\ & + \left(CF\right)\left(\delta^{Obj}_{DCF}\pi^{Obj} + \left(1-\delta^{Obj}_{DCF}\right)\pi_{t-1}\right) + \varepsilon_t,\end{aligned} \tag{8}$$

where $\pi^{e}_{t|t+n}$ is inflation expectations formed in period t for the forecast horizon $t+n$; π^{Obj} is the inflation target; π_{t-1} is lagged inflation; δ^{Obj} is the weight of the inflation target in expectations; and ε_t is the regression error.

Table 5

CREDIBILITY OF EXPECTATIONS WITH THE INFLATION TARGET

	δ^{Obj}	δ^{Obj}_{ACF}	δ^{Obj}_{DCF}	$H_0 : \delta^{Obj}_{ACF} = \delta^{Obj}_{DCF}$
Twelve-month ahead expectations	0.42[a]	0.36[a]	0.47[a]	−1.73
Four-year ahead expectations	0.66[a]	0.53[a]	0.70[a]	−3.99
Eight-year ahead expectations	0.76[a]	na	na	na

Note: Ordinary least square estimates were performed with Newey-West standard errors. [a] and [b] denote statistical significance at the 1% and 5% level, respectively. The value reported for the hypothesis test is the *t* statistic. na stands for not available.

Table 5 reveals that the coefficient δ^{Obj} is significant for all forecast horizons and increases as the horizon becomes longer. For short-term expectations, δ^{Obj} takes a value of 0.42, for medium-term ones this figure is 0.66, and for long-term ones it is 0.76. The outcomes clearly demonstrate that the anchoring of inflation expectations is influenced by the announcement of a long-run inflation target.

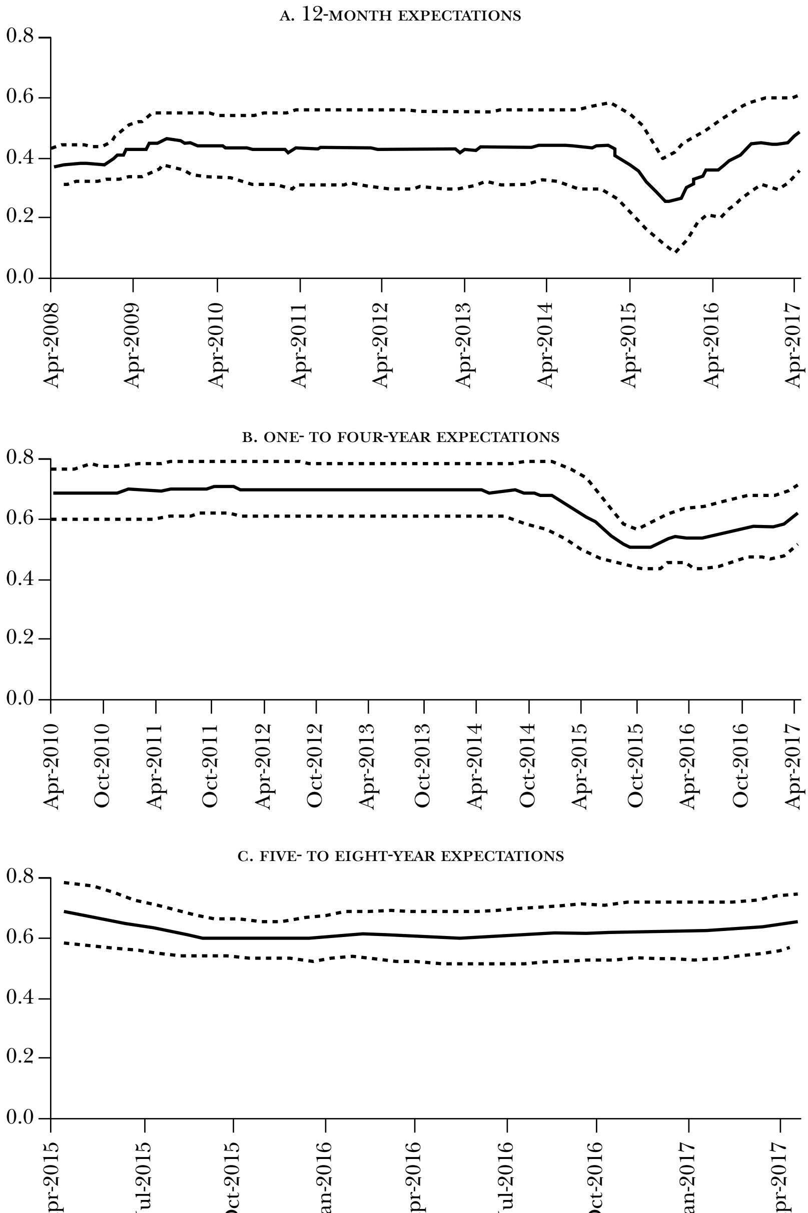
In addition to the above, it is important to underline that the coefficient δ^{Obj} displays an increase as compared to the value it showed before the 2008 crisis in short and medium-term expectations, which is mainly due to the communication tools used by the central bank during the last decade.

Figures 8a and 8b, employing six-year rolling regressions, reveal that the weight associated to the long-run target in short and medium-term inflation expectations has remained relatively stable most of the time, although it decreased at the start of 2015, possibly due to the volatile domestic and international economic environment. However, it is important to mention that said coefficient has returned to values similar

Figure 8

EXPECTATIONS CREDIBILITY WITH INFLATION TARGETING

Six-year rolling windows (coefficient δ)



to those registered before 2015 in both expectations. Meanwhile, Figure 8c shows that δ^{Obj} has remained unchanged for long-term expectations, which might be explained by the fact that long-term expectations are mainly determined based on the inflation target.

5.3.2 Credibility of Expectations, a VAR approach

This subsection follows the methodology of Demertzis et al. (2008, 2009) and uses a VAR model to assess the implicit anchoring of inflation expectations. In particular, long-term expectations are evaluated together with actual inflation. By being a VAR model, it attempts to explore the interdependence between both variables assuming that they are intrinsically related. The model seeks to measure the credibility of monetary policies given that if there is little correlation between the variables it would mean expectations are well-anchored. Due to the fact that a Cholesky decomposition is used to identify the model, the order of the variables is important. To maintain consistency with my earlier findings, in which expectations are not affected by contemporaneous inflation, the order employed in the VAR is to first specify the equation for inflation expectations followed by the equation for inflation. The selection of lags is carried out based on the Schwarz criterion. In specific, each model of 1 to 12 lags was evaluated, selecting the most parsimonious from them. The optimal number of lags is two for all the models. The generalization of the estimated model is as follows:

9

$$\pi^e_{t|t+n} = \gamma_0 + \gamma_1 \pi_{t-1} + \ldots + \gamma_p \pi_{t-p}$$
$$+ \theta_1 \pi^e_{t-1|t+n-1} + \ldots + \theta_p \pi^e_{t-p|t+n-p} + \varepsilon_{1t},$$

10

$$\pi_t = \alpha_0 + \alpha_1 \pi_{t-1} + \ldots + \alpha_p \pi_{t-p}$$
$$+ \beta_1 \pi^e_{t-1|t+n-1} + \ldots + \beta_p \pi^e_{t-p|t+n-p} + \varepsilon_{2t}.$$

The long-run solution to equations 9 and 10 takes the form:

$$\pi = \frac{\alpha_0}{1-\alpha_1-\ldots-\alpha_p} + \frac{\beta_1+\ldots+\beta_p}{1-\alpha_1-\ldots-\alpha_p}\pi^e,$$

$$\pi^e = \frac{\gamma_0}{1-\theta_1-\ldots-\theta_p} + \frac{\gamma_1+\ldots+\gamma_p}{1-\theta_1-\ldots-\theta_p}\pi.$$

The solutions to inflation and credibility are:

$$\lambda\pi^* = \frac{\gamma_0}{1-\theta_1-\ldots-\theta_p},$$

$$1-\lambda = \frac{\gamma_1+\ldots+\gamma_p}{1-\theta_1-\ldots-\theta_p}.$$

Simplifying and rearranging the expressions implies that:

$$\pi^* = \frac{\gamma_0}{1-\theta_1-\ldots-\theta_p-\gamma_1-\ldots-\gamma_p},$$

$$\lambda = 1-\frac{\gamma_1+\ldots+\gamma_p}{1-\theta_1-\ldots-\theta_p}.$$

Table 6 shows the implicit anchor for inflation expectations at the three horizons, revealing that for all of them the estimated value is relatively close to the long-run target of 3% set by Banco de México, the value being closest to 3% corresponding to long-term expectations.[7] Meanwhile, the weights of

[7] Outcomes for implicit inflation and creditability remain stable when the number of lags is changed.

implicit anchors of inflation expectations grow with respect to the horizon of the expectations. Thus, said value is 0.74 for short-term expectations, 0.92 for medium-term ones and 0.95 for long-term ones. The evidence therefore suggests that the relative importance of implicit anchor behavior increases as the forecasting time horizon becomes longer.

In addition to the above, analysis of pre- and postcrisis periods is performed, revealing that after the 2008 financial crisis the weight assigned to expectations increases at all forecast horizons, suggesting central bank credibility has grown over the last decade.

Table 6

IMPLICIT ANCHORING AND ITS CORRESPONDING WEIGHT IN THE FORMATION OF EXPECTATIONS

Sample		*Complete*	*Precrisis*	*Postcrisis*
Twelve-month ahead expectations	π^*	3.79	3.67	3.85
	λ	0.74	0.63	0.75
Four-year ahead expectations	π^*	3.52	3.42	3.52
	λ	0.92	0.74	0.95
Eight-year ahead expectations	π^*	3.4	na	na
	λ	0.95	na	na

Note: the number of optimal lags was obtained using the Schwarz criterion. The value reported for the hypothesis test is the *t* statistic. na stands for not available.

Figure 9 depicts the responses of short, medium and long-term expectations to an inflation shock of one standard deviation. The short-term response is statistically significant five months after the shock occurred and becomes nonsignificant eighteen months after it. The response of medium-term expectations is significant three months after the shock and its effect is not significant approximately one year after it. Finally, the response of long-term expectations to an inflation shock follows a similar path to that of medium-term ones, although to a lesser degree. In sum, the behavior of impulse responses at deferent horizons can be grouped into shocks that disperse faster and smaller impacts of inflation on expectations as the forecast horizon increases.

The speed with which impulse responses at different horizons become nonsignificant might be determined by the with which lag monetary policy operates; that is, after a shock, economic agents expect the central bank to act in a consistent manner to reduce its impact. The speed of adjustment would therefore depend on the persistence of inflation expectations in the face of different shocks, the structure of the economy, nominal and real rigidities, and the central bank's level of credibility among economic agents. Nevertheless, under a credible inflation targeting regime such as that in Mexico, shocks are expected to become diluted and impulse responses eventually converge to zero.

Figure 10 shows that the pre-2008 financial crisis vector autoregression exercise gives similar results to the exercise for the full sample, revealing that for medium-term expectations the shock dissipates in half the time taken for 12-month ahead expectations. Moreover, the size of the shock, in the same way as responses for the full sample, becomes smaller as the forecast horizon increases.

Performing the exercise for the postcrisis period it can be seen that for short-term expectations the period in which expectations respond to inflation decreases. The size of the shock is also smaller. Meanwhile, the response of medium-term expectations to an inflation shock is practically not significant for all the periods. The outcomes reflect a greater level of anchoring of expectations in the period after 2008 (Figure 11).

Figure 9

COMPLETE SAMPLE 2002M1-2017M3

Inflation and expectations response to a one standard deviation shock on inflation

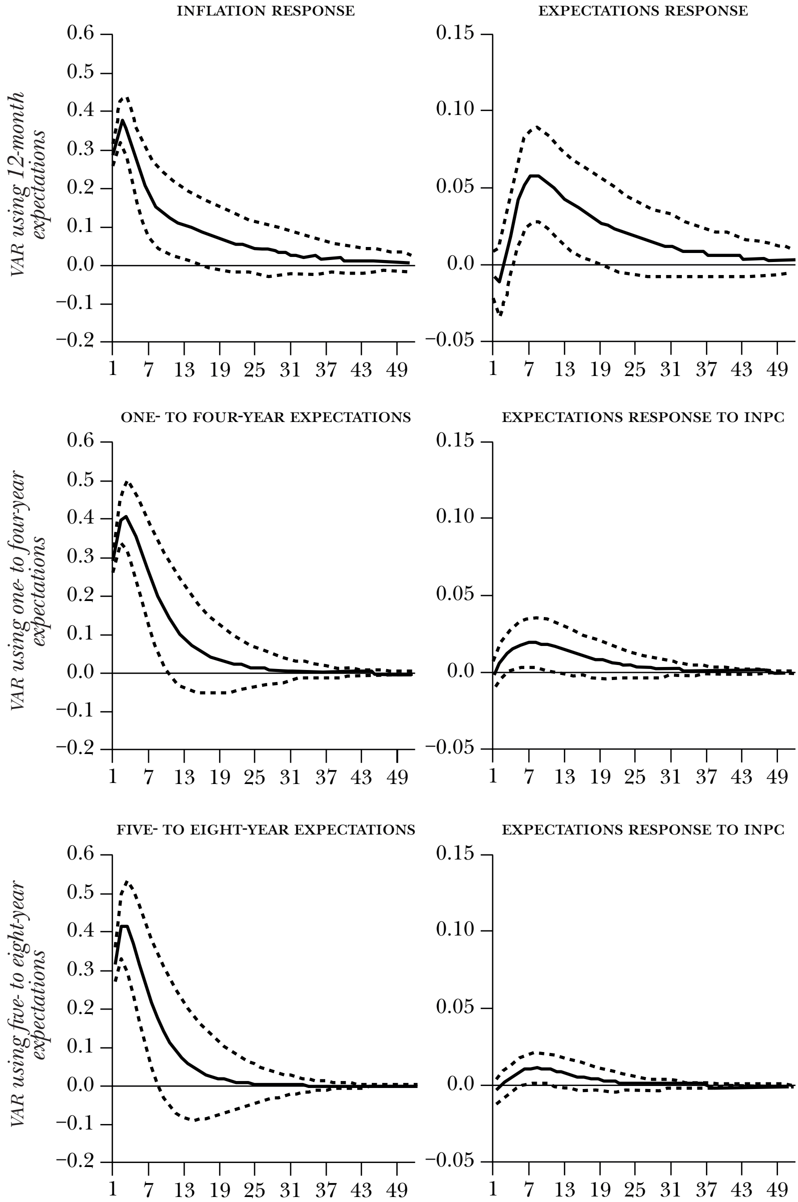


Figure 10

PRECRISIS SAMPLE 2002M1-2008M3

Inflation and expectations response to a one standard deviation shock on inflation

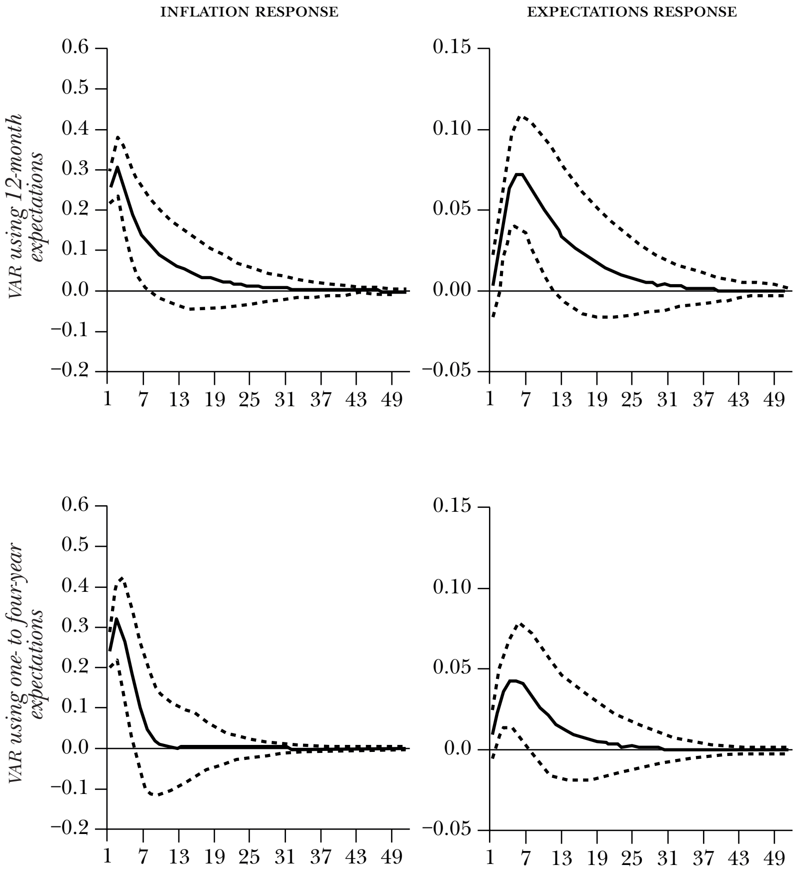
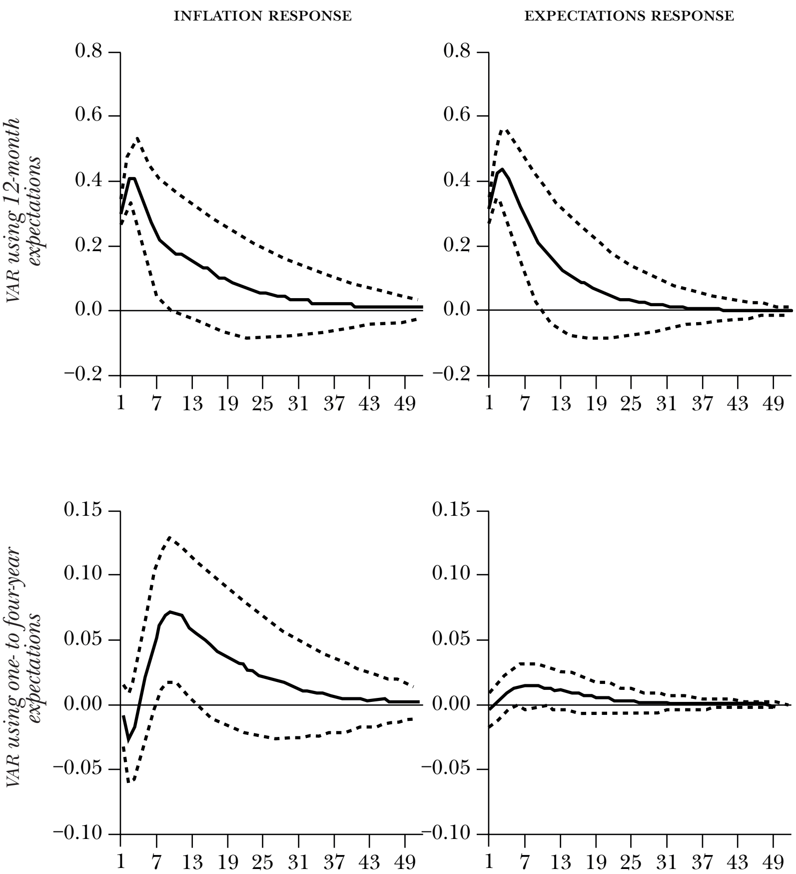


Figure 11

POSTCRISIS SAMPLE 2008M4-2017M3

Inflation and expectations response to a one standard deviation shock on inflation



6. CONCLUSION

This paper assessed the anchoring of inflation expectations introducing a novel classification according to the characteristic studied using econometric methods. In particular, three dimensions of anchorage were examined: sensitivity, resilience and credibility for the period between January 2002 and May 2017, as well as for two subsamples divided by the 2008 financial crisis.

The outcomes demonstrate that short-term expectations are more sensitive, followed by medium-term ones, while long-term ones are not affected by movements in inflation. They also highlight that after the 2008 financial crisis medium and long-term expectations are less sensitive to lagged inflation as well as short-term expectations.

Evidence was provided on how inflation shocks do not influence the formation of medium and long-term expectations, while short-term expectations are resilient to shocks after 2010 according to a moving windows analysis. Moreover, the credibility of Banco de México with regard to its long-run inflation target appears to have increased after the 2008 financial crisis despite substantial volatility in the markets.

It is evident that the analysis of anchoring using the dimensions of sensitivity, resilience and credibility not only facilitates study but also the reporting of outcomes. Nevertheless, this paper does not provide a guide on which of these dimensions is the most important with regards to deanchoring. For this reason, future efforts should focus on assessing the risks associated to each of those dimensions in order to reduce follow-up costs.

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Estimating and Forecasting Default Risk: Evidence from Jamaica

Andrene Senior
Sherene A. Bailey

Abstract

This paper employs the generalized method of moments estimation technique to evaluate the impact of macroeconomic factors on bank default risk for listed Jamaican banks and securities dealers over the period December 2004 to June 2016. Default risk is captured by a distance to default measure which is computed using a Merton type, option-based model. This indicator accurately tracks the default experience of listed Jamaican banks and securities dealers over important dates throughout the sample period. The estimation results of the model revealed that gross domestic product growth, inflation, unemployment rate, growth in domestic private sector credit as well as the real effective exchange rate have a statistically significant impact on the performance of the distance to default measure. As such, the econometric findings validate the sensitivity of the fragility measure to the variability of key macroeconomic variables. The model was also utilized to forecast the distance to default measure six-quarters ahead, as this will aid in the formulation of policy to mitigate systemic risks in the financial sector. The forecast results showed less volatility and lower overall default risk for Jamaican banks and securities dealers due to the projected improvement in various macroeconomic indicators.

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Keywords: default risk, forecasting, macroeconomic factors.
JEL classification: E17, F4.

1. INTRODUCTION

With more frequent instances of widespread distress during the last few decades, financial stability has become an increasingly important objective for policymakers. Episodes of profound banking system distress have occurred not only in emerging and developing economies but also in advanced industrialized countries, such as United States and Japan. In many cases, banking sector calamities have resulted in large losses of wealth and led to disturbances in the supply of credit within the economy. Furthermore, resolving these crises has frequently imposed a significant burden on public funds. These serious consequences underscore the value of indicators that signal a rising probability of banking sector problems before such problems actually occur and therefore represent an important aspect of effective banking supervision and financial market surveillance.

The approach to the development of measures of financial system distress has changed over the years and the locus of concern has shifted from examining solely microprudential indicators to also incorporating macroprudential dimensions of stability. Against this background, there has been increasing emphasis on early warning and forward-looking measures which can signal the risk of default of individual institutions as well as the system. These measures are useful in identifying the build-up of risks and potential vulnerabilities and would facilitate and enable a timelier reaction by the relevant authorities to any financial sector weaknesses which may arise. The distance to default is one such quantitative measure of financial stability which has been increasingly used by a number of central banks and international financial institutions. It is a widely used indicator of default risk and is a market-based risk measures for banks and nonfinancial corporates and captures the probability that the market value of a firm's assets falls below the value of its debt.[1] Market-based risk measures aim at supplementing more traditional analyses based on financial statements and income account statements with the added advantage of using the forward-looking information incorporated into security prices. Empirical studies have shown that the distance to default predicts well rating downgrades of banks in developed countries and emerging market countries. There is also empirical support for using the distance to default for financial institutions as a forecasting tool of bank distress.

Regarding Jamaica, based on a study by Lewis (2010), distance to default and the probability of default estimates were computed for the sovereign and for publicly listed financial institutions in the bank and nonbank sector in Jamaica for the period 2005 and 2010. The results underscored that these estimates serve as an early warning indicator of macrofinancial vulnerabilities during known periods of distress. Mingione (2011) also utilized principal component analysis to forecast indices of financial vulnerability for the Jamaican banking sector. He found that the principal component analysis model leads to more accurate predictions over the out-of-sample period using an aggregate index of vulnerability. Based on the literature, the forecast of these measures are useful in enabling policymakers and financial system participants to better monitor the degree of stability of the financial system as well as anticipate the sources and causes of financial stress to the system.

This paper builds on prior work for Jamaica by investigating the macroeconomic factors which impact banks' distance to default measures. The paper also provides a six-quarter ahead forecast of these institutions' distance to default using the generalized method of moments (GMM) estimation technique in order to gauge the degree of solvency and systemic risks within the banking sector. The paper is organized as follows: Section 2 provides an overview of the literature on the impact of macroeconomic factors on institutions' distance to default. In Section 3, there is a summary of the distance to default methodology

[1] See Tudela and Young (2003) and Chan-Lau (2006).

as well as trends in the measure for financial institutions listed on the Jamaica Stock Exchange. Section 4 provides a brief outline of the data used in the study as well as the estimation technique employed, while Section 5 presents the findings of the model. The conclusion and policy implications are presented in Section 6.

2. LITERATURE REVIEW

Bernoth and Pick (2009) forecasted systemic risk-taking into account linkages within the financial sector irrespective of whether they are caused by direct financial linkages or common shocks to the financial system. The study combined the use of unobserved common factors and observed variables for forecasting in a panel data set spanning 211 banks and 120 insurance companies in 21 countries. More specifically, it examined the importance of a number of macroeconomic variables and unobserved factors on the performance of banks and insurances. Against this background, there was an investigation of the forecast performance of macroeconomic and factor-augmented models of the fragility of banks and insurance companies. Also, given that the performance of firms in two industries and in geographically distinct regions was analyzed, there was an examination of the importance of regional, industry-specific or worldwide factors in forecasting financial fragility.

Furthermore, the study utilized distance to default as the measure of the performance of banks and insurance companies. It is based on the theoretical option pricing model of Merton (1974). An advantage of the distance to default is that it combines information about stock returns with leverage and volatility information and is, therefore, a more efficient indicator of default risk than simple equity price-based indicators.[2]

The explanatory variables included in the model are the growth rate of the 10-year bond yield, industrial production,

[2] See Vassalou and Xing (2004)

inflation, domestic credit, equity returns, real effective exchange rate, unemployment rate, price earnings ratio and the Chicago board of exchange volatility index. The results indicated that unobserved common factors play an important role, in particular taking unobserved factors into account leads up to 11% reduction in the root mean squared error (RMSE) of the forecasts of individual firms' distance to default. Systemic risk can also be better forecasted as the aggregate RMSE is reduced by 29% in one-quarter ahead forecasts and by 23% in four-quarter ahead forecasts.

Laurin and Martynenko (2009) quantitatively examined the relation between corporate default probability and macroeconomic information using panel data analysis. They also performed a quantitative comparison of default probability and macroeconomic information between different Swedish stock indexes based on market capitalization. The firms were segmented based on market capitalization. More specifically, a large-capitalization index was used, which consisted of firms with market capitalization of one billion euros, a mid-capitalization index included firms with market capitalization over 150 million euros but less than one billion euros and a small-capitalization index comprising firms with capitalization up to 150 million euros. The explanatory variables used were the domestic industrial production index, consumer price index, nominal domestic three-month rate for Treasury bills (R3M), GDP-growth, unemployment rate, exchange rate, equity price index and a measure of equity volatility. An autoregressive model with one-year lagged distance to default is also estimated. [3]

[3] Autoregressive models are often used in studies of time series data where the behaviour of a dependent variable is determined by its previous estimations. Åsberg and Shahnazarian (2008) presented an estimation model for predicting the distance to default. The model is based on the hypothesis that the best forecast for future distance to default is provided by the recent outcomes for the variable in question.

The panel regression results for the large-capitalization and the mid and small-capitalization firms appeared to be similar. It was found that the one-year lagged Industrial Production Index and the one-year lagged exchange rate exhibited a large negative effect on the probability of default. The interest rate and the one-year lagged interest rate were found to have a positive impact on the probability of default. The autoregressive model, with an autoregressive lagged term, showed a decreasing distance to default over time.

In concluding, macroeconomic factors such as the one-year lagged industrial production index, the one-year lagged exchange rate, and the one-year lagged interest rate explained 75% of the changes in the probability of default for the large-capitalization firms (68% in the model for the mid- and small capitalization firms, respectively). The autoregressive model indicates a weak explanatory power and an increasing probability of default overtime.

Hamerle et al. (2004) forecasted credit default risk in loan portfolios using a Merton-style threshold-value model for the default probability which treats the asset value of a firm as unknown and where default correlations are also modeled. The empirical analysis is based on a large data set of German firms provided by Deutsche Bundesbank for the period 1987 to 2000. The data was collected by Deutsche Bundesbank's branch offices in order to evaluate the credit quality of firms for refinancing purposes.

Of importance, the inclusion of variables which are correlated with the business cycle improved the forecasts of default probabilities. Further, the better the point-in-time calibration of the estimated default probabilities, the smaller the estimated correlations, as such, correlations and default probabilities should always be estimated simultaneously. The macroeconomic variables included in the model were the business climate index, unemployment rate and systematic growth in new orders of the construction industry. The model allowed default probabilities to be forecasted for individual borrowers and estimated correlations between those borrowers simultaneously.

3. METHODOLOGY

3.1 Distance to Default Framework

The distance to default measure captures the probability that the market value of a firm's assets falls below the value of its debt. More specifically, the face value of debt is typically computed from balance sheet data and is assumed equal to the sum of the short-term liabilities plus half the long-term liabilities. The distance to default is then derived using the market value of the firm as well as the implied equity price volatility.

Distance-to-default is based on the structural model of corporate debt first introduced by Black and Scholes (1973) and Merton (1974). Furthermore, the framework is premised on the relation between the value of the firm, V_A (or the value of its assets), which should be equal to the sum of the values of its debt, X, and equity, V_E. In addition, typically the firm's assets are first used to pay debtholders while whatever is left is distributed to shareholders. In particular, the value of equity is shown in Equation 1:

$$V_E = \max(0, V_A - X). \quad (1)$$

Also, compensation to equity holders is equivalent to a call option on the value of the firm with a strike price equal to the face value of debt. The strike price is also known as the default barrier is set equal to the level of the firm's short-term liabilities and half its long-term liabilities. Information on the value of the firm, the debt owed by the firm and the market value of equity is enough to derive the remaining unknown variable.

According to the Black-Scholes (1973) model, the market value of the firm's underlying assets is due to the following stochastic process:

2 $$dV_A = \mu V_A dt + \sigma_A V_A dz,$$

where V_A and dV_A are the firm's asset value and the change in asset value; μ and σ_A are the firm's asset value drift rate and the volatility; and dz is a Wiener process.

Furthermore, according to the Black and Scholes (1973) and Merton (1974) option pricing theory, the equity call option written by debt holders to shareholders may be valued by solving the following second-order linear partial differential equation:

$$\frac{\partial V_E}{\partial t} = rV_E - rV_A \frac{\partial V_E}{\partial V_A} - \frac{1}{2}\sigma^2 V_A^2 \frac{\partial^2 V_E}{\partial V_A^2},$$

subject to the boundary conditions:

$$V_E\left(V_A,\ t\right)\begin{cases} = V_A - X, & V_A \geq X \\ = 0, & V_A < X \end{cases}$$

The unique solution to this partial differential equation is the celebrated Black-Scholes-Merton option pricing formula:

3 $$V_E = V_A N\left(d1\right) - e^{-rT} XN\left(d2\right),$$

where V_E is the market value of the firm's equity, $N\left(d\right)$ is the cumulative normal density function, and r is the risk-free interest rate. Solving Equation 3 for *d1* and *d2* yields the following expressions:

4 $$d1 = \frac{ln\left(\frac{V_A}{X}\right) + \left(r + \frac{\sigma_A^2}{2}\right)T}{\sigma_A \sqrt{T}},$$

5 $$d2 = \frac{ln\left(\frac{V_A}{X}\right) + \left(r - \frac{\sigma_A^2}{2}\right)T}{\sigma_A \sqrt{T}} = d1 - \sigma_A \sqrt{T}.$$

Of note, $d2$ shown in Equation 5 represents the distance to default, where $\left(V_A/X\right)$ captures the firm value relative to the default threshold, which over time is impacted by the interest rate and asset value volatility. This distance to default expression is then standardized by the volatility of the firm's assets.

3.2 Trends in Distance to Default for Financial Institutions Listed on the Jamaica Stock Exchange

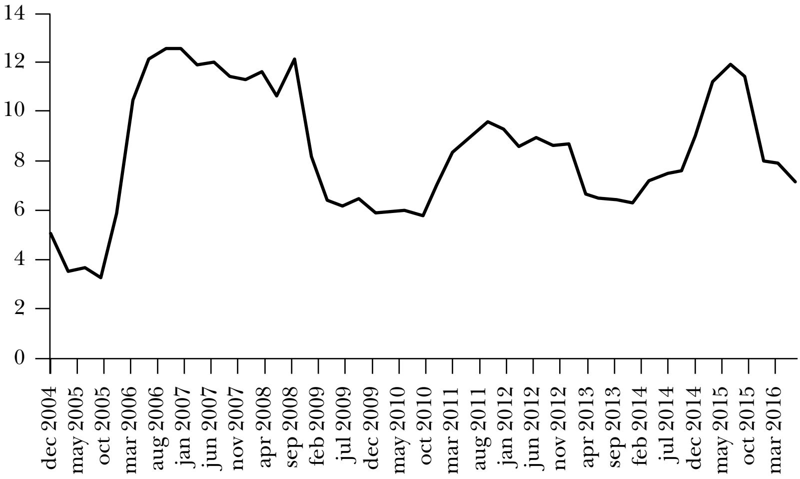
The distance to default was successful in tracking the default experience of listed banks during periods of vulnerability throughout the sample period (see Figure 1). The measure declined during the global crisis period, indicating that there was deterioration in the default measure of these institutions during this period. This occurred in a context where the crisis would have contributed to declines in the value of the asset holdings of these institutions. In addition, the measure also fell during the two debt exchange periods in Jamaica, which occurred in 2010 and 2013 and which involved the extension of maturity and reduction of coupon rates on local currency denominated Government of Jamaica bonds.[4] The distance to default measure was adversely impacted by weaker profitability performance of the listed banks due to the lower revenue performance on these investments.

The distance to default for the securities dealers declined or remained low throughout periods of vulnerability, such as during the two debt exchanges which occurred during 2010 and 2013 (see Figure 2). The measure was adversely impacted by weaker profitability performance of the listed securities dealers due to the lower revenue performance on domestic currency Government of Jamaica investments. Securities dealers have also been impacted by the continued phasing down of the

[4] The Jamaica Debt Exchange occurred in the March 2010 quarter and the National Debt Exchange took place during the March 2013 quarter.

Figure 1

DISTANCE TO DEFAULT: DEPOSIT-TAKING INSTITUTIONS LISTED ON THE JAMAICA STOCK EXCHANGE

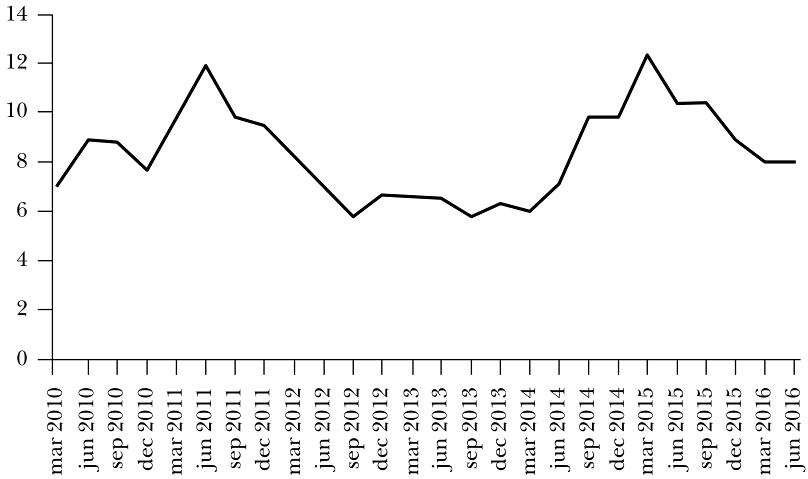


retail repurchase business of the sector since 2015.[5] This has coincided with weaker profitability and lower distance to default values for these institutions during this period.

[5] Securities dealers' fund the purchase of securities through repurchase agreements (repos). The risks embedded in these repos emanate from securities dealers' reliance on borrowing very short-term funds from retail clients and institutional investors to take proprietary positions in primarily long-term government securities. To address the systemic risks from these broker-dealer activities, the Government of Jamaica committed to reform the broker-dealer industry, which included the phasedown of the *retail repo* business model. Legislation was enacted to allow for the establishment of the Collective Investment Scheme, which facilitates the transfer of market, interest rate and liquidity risk to individual investors and off the balance sheet of broker dealers. As a result, since 2013, the securities dealers' sector embarked on a process of reform which entailed the phasedown of the *retail repo* business model.

Figure 2

DISTANCE TO DEFAULT: SECURITIES DEALERS LISTED ON THE JAMAICA STOCK EXCHANGE



4. EMPIRICAL ANALYSIS

4.1 Data and GMM Estimation Technique

The paper employs quarterly distance to default data for banks and securities dealers listed on the Jamaica Stock Exchange as well as information on selected macroeconomic variables over the period December 2004 to September 2016. Macroeconomic variables utilized in the study included nominal GDP growth, growth in the inflation and unemployment rates, growth in the real effective exchange rate (REER), changes in the 10-year GOJ global bond yields, growth in private sector credit, and the spread between loan and time deposit rates.

Panel data estimation was used as it facilitates the inclusion of time series data across several variables. Panel data analysis also makes it possible to predict the behavior of the individual variables more precisely than other techniques as it utilizes time series data and therefore captures the past experiences of each variable. More specifically, the GMM estimation technique was employed to estimate the relation between distance to default and macroeconomic variables for both banks and securities dealers.[6] The technique was chosen as it uses assumptions about specific moments of the random variables instead of assumptions about the entire distribution. The GMM method is also useful in providing unbiased and efficient estimates in dynamic models which have lagged endogenous variables as regressors. Based on work by Boucinha and Ribeiro (2007), the methodology can be utilized to obtain consistent estimates of the parameters of interest when the persistence of the dependent variable needs to be modelled explicitly. Furthermore, the model does not require strong hypotheses about the exogeneity of the regressors. Arellano and Bond (1991) suggest that consistent and efficient estimates can be obtained by using lagged values of the dependent variable and lagged values of the exogenous variables as instruments. Baltagi (2001), also highlighted that the GMM methodology accounts for the possibility of correlations between the independent variables, making it an advantageous technique.

More specifically, the GMM estimation technique shows how a variable in period t, for example, y_{it}, could be explained through the value of the same variable in period $t-1$, $y_{i,t-1}$, along with other different explanatory elements, x'_{it}, and a random error term, η_{it}. This relation is outlined in Equation 6.

$$y_{it} = \alpha + \delta y_{i,t-1} + x'_{it}\beta + \eta_{it}, \tag{6}$$

where y_{it} is the dependent variable, α is the intercept, δ is a scalar, β is the $k \times 1$ vector of explanatory variables' parameters, x_{it} is the $1 \times k$ vector of explanatory variables, with Equation 7 explaining the random error term, η_{it} which includes

[6] Of importance is that the bond yield variable was only included in the model for the securities dealers.

individual unobserved effects, μ_i, and the genuine random error term, ε_{it}.

$$\eta_{it} = \mu_i + \varepsilon_{it} , \tag{7}$$

where $\mu_i \sim \text{iid}\left(0, \sigma_\mu^2\right)$ and $\varepsilon_{it} \sim \text{iid}\left(0, \sigma_\mu^2\right)$ are independent of each other and themselves.

Furthermore, concerning the matter of autocorrelation as it relates to the GMM framework, Arellano and Bond (1991) utilized internal instruments that are lagged values of the levels of the variables which appear on the right-hand side of Equation 6 in addressing this issue. These instrumental variables should not be correlated with the first difference of the error term but should be correlated with the variable to be estimated. The idea behind this technique is to estimate the model by combining several instruments around a single vector of parameters, in order to obtain the minimum correlations between the error term and the relevant instruments. In particular, this technique considers as suitable instruments of the second- and higher-order lags of the regressors in the event of no serial correlation in the time-varying component of the disturbance term.

5. RESULTS

5.1 GMM Model

Panel unit root tests were done on the residuals of the GMM model for each sector. More specifically, the unit root tests applied were the Levin, Lin and Chu test, Im, Peasaran and Shin test, ADF-Fisher Chi-square test and PP-Fisher Chi-square test. All the tests showed that the residuals for both models are stationary, reflecting a nonspurious regression (see Tables A.3 and A.6). Additionally, the Sargan test of orthogonality between the instruments and the residuals, which tests the validity of instruments used in the regression through a comparison between the estimated moments and the sample moments was used to evaluate the results. The Sargan test results showed that there was no evidence to reject the null that "over-identifying restrictions are valid," which suggests that the instruments used in the models are valid.

5.1.1 DTI Results

The results of the GMM model were consistent with expectations. All macroeconomic variables included in the model, with the exception of the growth in the REER index, have a statistically significant impact on the distance to default measure. In particular, the findings showed a positive relation between GDP growth and the distance to default. Stronger performance in GDP growth is expected to contribute to stronger bank performance, for instance through increased deposit growth and investments, which will ultimately lead to improvements in these institutions' distance to default. There is also a positive relation between the loan rate and time rate deposit spread and the distance to default. An increase in this spread typically contributes to improvement in the revenue performance of banks and should lead to increases in the distance to default.

An increase in the growth of the unemployment rate resulted in deterioration in the distance to default. This is anticipated given that worsening in the unemployment rate is expected to increase nonperforming loans of banks and worsen performance. Based on the literature, the relation between growth in domestic credit to the private sector and financial institution performance is ambiguous. Some studies, such as Hagen and Ho (2004) and Goldstein (1998), indicate that there is a negative relation between credit growth and distance to default, as banking distress is typically preceded by credit booms.[7]

[7] Work by Bernoth and Pick (2009) showed a positive relation between credit growth and distance to default, indicative of stronger credit growth improving the profitability of banking institutions.

The findings of this study also show an inverse relation between growth in private sector credit and distance to default. Furthermore, stronger growth in inflation was also found to negatively impact distance to default, as deterioration in inflation performance can tend to erode the profitability of banking institutions. Additionally, the lagged dependent variable was positive and statistically significant, and is indicative of the persistence of the dependent variable in explaining itself.

The model has a high R-squared of 76.1% and a Durbin-Watson statistic of close to two. Furthermore, period dummies for the global crisis period and the National Debt Exchange period were found to be significant.

5.1.2 Forecast Performance and Forecast Evaluation Results

The results of the GMM model in Section 3.1 were used to generate both in-sample and out-sample forecasts of the distance to default measure. The in-sample estimates were generated over the entire sample period, March 2004 to June 2016, while the out-of-sample estimates were generated for the period, December 2014 to June 2016. The summary statistics for these estimations are reported in Table A.1 and Table A.2

The forecasting ability of the GMM model was evaluated using common measures such as the Theil inequality coefficient (Theil U) statistic and the root mean square error (RSME). The Theil U statistic is useful is determining a model's prediction performance relative to a naïve model, which is a benchmark used for evaluating forecast accuracy where the forecast assumes that the value in the next period is the same as the value in this period. Furthermore, the Theil U coefficient lies between zero and one; with values closer to zero, indicative of greater accuracy of the prediction model. Additionally, the root mean squared error is calculated based on the square root of the squared difference between predicted and observed values, where lower values are indicative of better forecasting ability of the model.

The prediction performance of the model was assessed using in-sample and out-of-sample forecasts. In-sample performance statistics based on the Theil U and RSME were 0.2 and 3.3, respectively, while the respective values for the out-of-sample forecast were 0.1 and 2.7. These results confirm that the model utilized has strong predictive power.

Given the strong predictive power of the model, which relied on projections of specific macroeconomic variables, the model was used to project the distance of default of listed deposit-taking institutions (DTIs) up to December 2017. For the banking sector, the findings showed that growth in the inflation rate, growth in private sector credit, bank spreads, growth in the unemployment rate and GDP had a statistically significant impact on the distance to default of these institutions. Of note, the unemployment rate, growth in private sector credit and growth in inflation have an inverse relation with DTIs' distance to default. The forecast for the distance to default of the banking sector was generally low and also reflected much lower volatility. This forecasted performance is largely due to the projected orderly movements of the statistically significant macroeconomic variables, in particular, credit growth and the unemployment rate.

5.1.3 Securities Dealers Results

Consistent with expectations, the finding showed a significant inverse relation between the distance to default and growth in the inflation rate. Similar to the DTIs, deterioration in this predictive variable is expected to have an adverse impact on the distance to default as deterioration in inflation performance can lead to higher expenses for the financial institutions and weaken profitability. The results also indicate a significant inverse relation between the distance to default and growth in private sector credit, as it is often the case that financial system fragility is sometimes preceded by marked acceleration in credit growth. Unlike for the DTIs, it was found that there is a significant inverse relation between the distance to default and GDP growth. This performance may occur because stronger performance in GDP growth may lead to higher funding demand, increased interest costs, higher bond yields and lower bond prices, which will ultimately lead to deterioration in these institutions' distance to default. There is also a positive relation between the loan and time deposit rate spread and the distance to default. An increase in this spread typically contributes to improvement in the revenue performance of banks and should lead to increases in the distance to default.

The results also showed that the growth in the REER index, return on GOJ global bonds and growth in the unemployment rate do not have a statistically significant impact on the distance to default. Nonetheless, as in the case of the DTIs, the lagged dependent variable was positive and statistically significant and is also indicative of the persistence of the dependent variable in explaining its own performance.

The R-squared of the model is 62.8%, and it suggests that the variables employed have a strong impact in explaining the performance of the distance to default. Additionally, period dummies for the National Debt Exchange period as well as the dummy capturing the periods of reform as it relates to the securities dealers' business model were found to be significant.

5.1.4 Forecast Performance and Forecast Evaluation Results

Based on the GMM model in Section 3.1, an in-sample forecast of the distance to default measure was done for the entire sample period, March 2010 to March 2016, while the out-of-sample forecast covered the period from March 2015 to March 2016. The in-sample performance statistics based on the Theil U and RSME were 0.1 and 2.0, respectively, while the respective values for the out-of-sample forecast were 0.08 and 0.8. The results also confirmed the strong predictive power of this model.

This GMM estimation techniques was also used to project the distance of default for the SDs' sector up to December 2017. For the SDs' sector, growth in the inflation rate, private sector credit growth, GDP growth and banks' interest rate spreads had a statistically significant impact on the distance to default of these institutions. Of note, growth in inflation has a negative relation with SDs' distance to default. The forecast for the distance to default of the SDs' sector also reflected lower volatility. This forecasted performance is largely due to the projected orderly movements of the statistically significant macroeconomic variables, in particular, credit growth and GDP.

6. CONCLUSION AND POLICY IMPLICATIONS

The distance to default measure utilized in the study was useful in identifying important dates throughout the sample period, where financial institutions would have experienced increased likelihood of insolvency. The periods included the recent global crisis period and the Jamaica Debt Exchange and National Debt Exchange periods during 2010 and 2013, respectively.

In addition, the GMM estimation technique was also used to determine the impact of macroeconomic factors on the distance to default of DTIs and SDs. For DTIs, the findings showed that growth in the inflation rate, growth in private sector credit, banks spreads, growth in the unemployment rate and GDP had a statistically significant impact distance to default of these institutions. Regarding the securities dealers, similar macroeconomic factors were found to impact default risk. In particular, the growth in the inflation rate, GDP, and the interest rate spread between loan rates and deposit rates had a significant impact on the distance to default.

The models were also used to forecast the distance to default, six quarters ahead, for both the DTIs and the SDs. Forecast results will be a useful tool in predicting the likelihood of financial institution distress and incorporates investors' forward-looking expectations. Findings for both DTIs and SDs showed trend improvement for the forecast period as well as significant reduction in volatility for the projected distance to default. The performance in the distance to default measure for the DTIs largely reflects the movement in GDP growth rate, inflation rate and the interest rate spread variable. For the SDs, forecast results were also largely underpinned by the performance of the inflation, GDP and interest rate spreads.

The findings re-emphasize the importance of consistency between Jamaica's macroeconomic program, which includes medium-term projections of the real, fiscal, external and monetary sectors, and the solvency of the banking sector. The forecast model is also useful in examining how severe movements in macro variables will impact the likelihood of institution failure. Furthermore, closer attention to market-based signals of risk, such as the distance to default, can enable regulators to be more proactive in implementing measures to limit the likelihood of a crisis or minimize its impact.

Distance to default forecasts can also be used as a forward-looking analytical tool to monitor systemic risk in the Jamaican financial system. Information contained in these forecasts can provide guidance for macroprudential policymakers, by signaling whether there is a build-up of systemic risks. This can fuel an evaluation by the relevant authorities as to the nature these vulnerabilities and whether the implementation of macroprudential tools are necessary to limit these risks.

Institution by institution findings can be useful in complementing work on systemically important financial institutions (SIFIs) by highlighting which of these institutions have a high degree of vulnerability to default risk. This is critical given that these institutions have a high degree of complexity and close linkages to the rest of the financial system and can pose a high risk to stability. Early signals of distress as it relates to SIFIs can aid in establishing a regulatory framework that can cope with risks arising from systemic linkages.

ANNEX. TABLES AND FIGURES

Table A.1

ESTIMATION OUTPUT FOR DEPOSIT-TAKING INSTITUTIONS DISTANCE TO DEFAULT

Sample (adjusted): 2005Q2-2016Q2
Periods included: 45
Cross-sections included: 2
Total panel (balanced) observations: 90
Instrument specification: GDP growth, inflation growth, spread, @sysper
Constant added to instrument list

Variable	*Coefficient*	*t-statistic*
Distance (−1)	0.917959	33.95348
GDP growth	12.42028	2.440430
REER growth (−2)	4.089674	1.899280
Credit growth	−7.395536	−3.279189
Inflation growth	−1.018786	−3.727524
Unemployment rate	−7.512652	−4.014348
Spread	0.075410	5.643401
@isperiod ("december2008")	−3.912005	−4.593268
@isperiod ("december2009")	0.146271	0.177412
@isperiod ("december2012")	0.348913	0.720158
@isperiod ("december2013")	−1.465139	−4.992281
@isperiod ("december2014")	0.598372	1.669097
Effects specification		
R^2		0.761039
J-statistic		29.61345
Durbin-Watson statistic		1.669466
Instrument rank		45

Table A.2

ESTIMATION OUTPUT FOR DEPOSIT-TAKING INSTITUTIONS' DISTANCE TO DEFAULT OUT-OF-SAMPLE FORECAST

Sample (adjusted): 2005Q2-2014Q4

Periods included: 45

Cross-sections included: 2

Total panel (balanced) observations: 78

Instrument specification: GDP growth, inflation growth, spread, @sysper

Constant added to instrument list

Variable	*Coefficient*	*t-statistic*
Distance (−1)	0.991793	21.62627
GDP growth	18.63147	2.950299
REER growth (−2)	2.121872	0.738073
Credit growth	−10.17660	−3.094955
Inflation growth	−0.390902	−1.768780
Unemployment rate	−7.244699	−2.959229
Spread	0.044987	1.895589
@isperiod ("december2008")	−4.057752	−4.755763
@isperiod ("december2009")	−0.393300	−0.404295
@isperiod ("december2012")	0.002545	0.005933
@isperiod ("december2013")	−1.670782	−3.909016
@isperiod ("december2014")	0.311304	0.767930
Effects specification		
R^2		0.761056
J-statistic		22.80316
Durbin-Watson statistic		1.707767
Instrument rank		39

Table A.3

DEPOSIT-TAKING INSTITUTIONS' DISTANCE TO DEFAULT ESTIMATION

Unit Root Results for the Residual

Sample: 2004Q1-2017Q4

Exogenous variables: individual effects

Balanced observations for each test

Method	*Statistic*	*Probability* [2]	*Cross-sections*	*Observations*
Null: Unit root (assumes common unit root process)				
Levin, Lin and Chu t^1	−7.73331	0.0000	2	88
Null: Unit root (assumes individual unit root process)				
Im, Pesaran and Shin W-stat	−6.37522	0.0000	2	88
ADF-Fisher χ^2	40.7064	0.0000	2	88
PP-Fisher χ^2	40.1889	0.0000	2	88

Note: [1]Under the null hypothesis, the test statistic is asymptotically disturbed according to the standard normal distribution. [2] Probabilities for Fisher tests are computed using an asymptotic χ^2 distribution. All other tests assume asymptotic normality.

Table A.4

ESTIMATION OUTPUT FOR SECURITIES DEALERS' DISTANCE TO DEFAULT

Sample (adjusted): 2010Q2-2016Q2

Periods included: 25

Cross-sections included: 4

Total panel (balanced) observations: 100

Instrument specification: @sysper, GDP growth, GOJ global bonds, spread, inflation growth, credit growth

Constant added to instrument list

Variable	*Coefficient*	*t-statistic*
Distance (−1)	0.408153	3.514498
Credit growth	−25.24730	−2.330699
GDP growth	−24.39533	−2.026492
Inflation growth (−1)	−1.117643	−2.454584
REER growth (−1)	−0.312925	−0.028075
GOJ global bonds	−0.203448	−0.800967
Spread	0.514586	4.153419
Unemployment rate	−1.848043	−0.426725
Constant	−1.162222	−0.501724
@isperiod ("december2011")	2.091702	2.850433
@isperiod ("december2013")	1.632662	1.994374
@isperiod ("december2014")	3.429162	3.756840
@isperiod ("december2015")	−0.512038	−0.796161

Effects specification

R^2	0.627477
J-statistic	16.33019
Durbin-Watson statistic	1.332565
Instrument rank	25

Table A.5

ESTIMATION OUTPUT FOR SECURITIES DEALERS' DISTANCE TO DEFAULT OUT-OF-SAMPLE FORECAST

Sample (adjusted): 2010Q2 2015Q4

Periods included: 23

Cross-sections included: 4

Total panel (balanced) observations: 92

instrument specification: @sysper, GDP growth, GOJ global bonds, spread, inflation growth, credit growth

Constant added to instrument list

Variable	*Coefficient*	*t-statistic*
Distance (−1)	0.548918	4.969056
Credit growth	−29.87750	−2.543776
GDP growth	7.064194	0.479160
Inflation growth (−1)	2.198643	1.821364
REER growth (−1)	−3.774137	−0.357726
GOJ global bonds	−0.833715	−2.526563
Spread	0.346364	2.418892
Unemployment rate	−2.697641	−0.582113
Constant	5.287464	1.677455
@isperiod ("december2011")	0.712883	0.847213
@isperiod ("december2013")	0.144404	0.152958
@isperiod ("december2014")	0.691408	0.574013
@isperiod ("december2015")	−0.436064	−0.591774
	Effects specification	
R^2		0.661071
J-statistic		13.59101
Durbin-Watson statistic		1.556667
Instrument rank		23

Table A.6

SECURITIES DEALERS DISTANCE TO DEFAULT ESTIMATION-UNIT ROOT RESULTS FOR THE RESIDUAL

Sample: 2010Q1 2017Q4

Exogenous variables: Individual effects

Balanced observations for each test

Method	*Statistic*	*Probability*[2]	*Cross-sections*	*Observations*
Null: Unit root (assumes common unit root process)				
Levin, Lin and Chu t[1]	−3.65842	0.0001	4	96
Null: Unit root (assumes individual unit root process)				
Im, Pesaran and Shin W-stat	−4.68516	0.0000	4	96
ADF-Fisher	35.2462	0.0000	4	96
PP-Fisher	35.4061	0.0000	4	96

Note: [1]Under the null hypothesis, the test statistic is asymptotically disturbed according to the standard normal distribution. [2]Probabilities for Fisher tests are computed using an asymptotic distribution. All other tests assume asymptotic normality.

Table A.7

GMM ESTIMATION OF DEPOSIT-TAKING INSTITUTIONS' DISTANCE OF DEFAULT

Forecast Performance Results

	In-sample forecast	*Out-of-sample forecast*	*Projections*
Forecast sample	2005Q2 to 2016Q2	2015Q2 to 2016Q2	2016Q2 to 2017Q4
Root mean squared error	3.33	2.66	1.00
Mean absolute error	2.58	2.05	0.82
Theil inequality coefficient	0.21	0.14	0.06

Table A.8

GMM ESTIMATION OF SECURITIES DEALERS' DISTANCE OF DEFAULT
Forecast Performance Results

	In-sample forecast	*Out-of-sample forecast*	*Projections*
Forecast sample	2010Q2 to 2016Q2	2015Q2 to 2016Q2	2016Q2 to 2017Q4
Root mean squared error	2.04	0.76	0.95
Mean absolute error	1.48	0.58	0.85
Theil inequality coefficient	0.14	0.08	0.09

Figure A.1

DTIS' ACTUAL, FITTED, RESIDUAL GRAPH

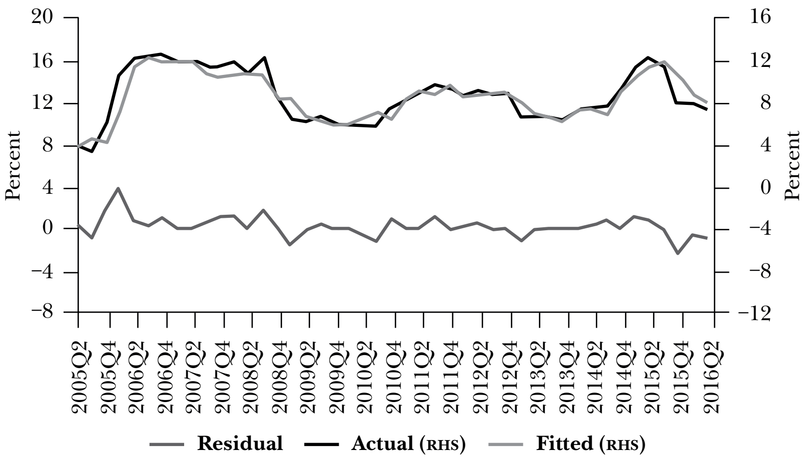


Figure A.2

SECURITIES DEALERS' ACTUAL, FITTED, RESIDUAL GRAPH

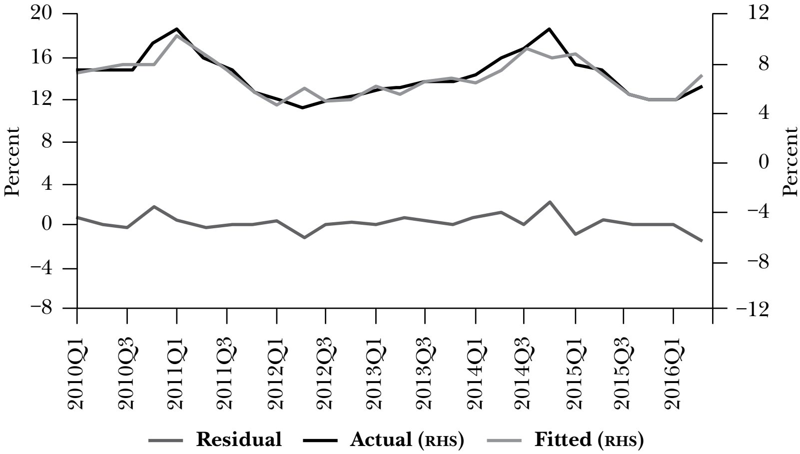


Figure A.3

DTIS' DISTANCE TO DEFAULT

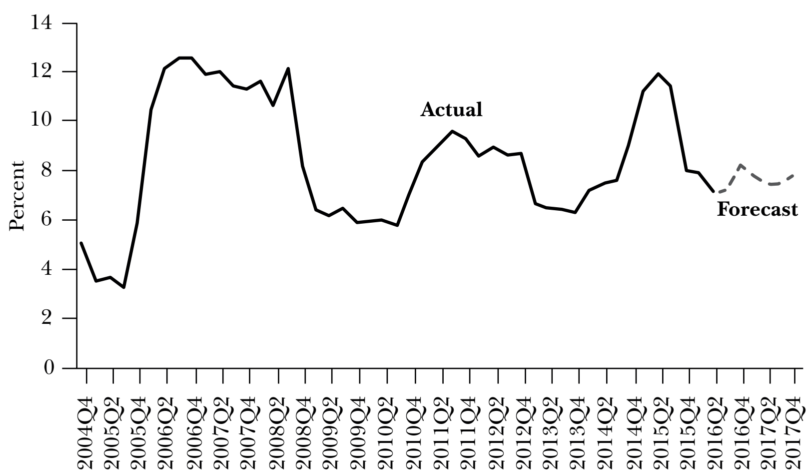
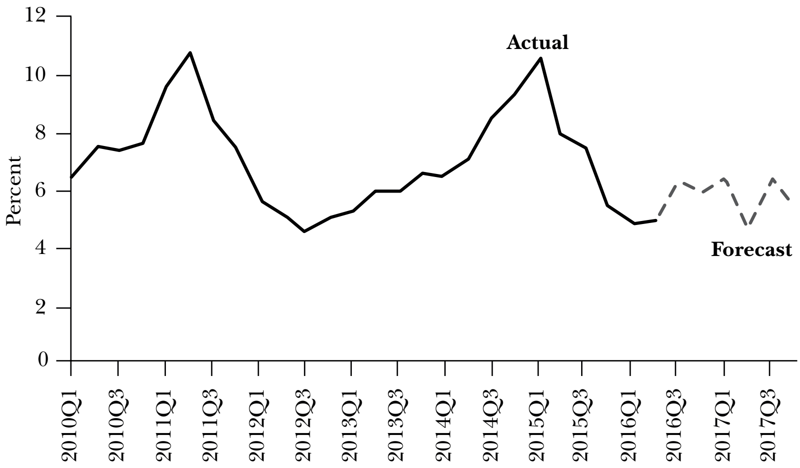


Figure A.4

SECURITIES DEALERS' DISTANCE TO DEFAULT



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