AN ECONOMETRIC ANALYSIS ON SURVEY-DATA-BASED ANCHORING OF INFLATION EXPECTATIONS IN CHILE*

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I. INTRODUCTION

To what extent are inflation expectations of public agents anchored to the Central Bank of Chile's (CBC) inflation target? In this note, I perform several econometric testing procedures in an attempt to answer this question. Expectations' anchoring is understood as another central bank instrument, in the sense that it helps to promote price stability through the ability of policy makers to influence inflation expectations. Thus, expectations anchoring comes out as a result of policy actions. I provide evidence suggesting that agents' expectations are indeed anchored. This is found within the manoeuvring horizon of monetary policy. Furthermore, within the fully-fledged inflation targeting period—starting in 2000—three periods are detected, namely (i) the pre-Global Financial Crisis (GFC) period, (ii) the post-GFC period until mid-2015, and (iii) the low inflation period afterwards. Particularly for the last period, the anchoring effect is clearer and more pronounced.

The econometric analysis relies on the Łyziak and Paloviita (2017) approach, which can be understood as a battery of estimates and statistical inference, plus some other robustness estimations found in the related literature. These estimations include the analysis proposed by Ehrmann (2015) on the dependence of inflation expectations on actual inflation, and the pass-through of shorter- to longer-term expectations. I also investigate anchoring through the Bomfim and Rudebusch (2000) method, provided the relationship between long-term expectations deviations and target deviations. This analysis is extended to consider two communication tools employed by the CBC, i.e. the inflation target and the inflation forecasts contained in its Monetary Policy Report ("IPoM", from its Spanish name *Informe de Política Monetaria*). Lastly, I consider the vector autoregression (VAR)-based analysis proposed by Demertzis et al. (2008, 2009) to derive the implicit inflation target contained in the expectations and the anchoring degree to that implicit target. For robustness purposes, I also analyse its dynamic evolution through a time-varying parameters version. I make use of the inflation expectations provided by the *Survey of Professional*

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Forecasters-equivalent for Chile[1], which despite the traditional criticism deserved for survey-based data contain several statistical advantages.[2] A special treatment to the GFC is given, as represent the period with major turbulences and de-anchoring evidence.

The anchoring of inflation expectations must be understood within the boundaries of the modern theory of optimal monetary policy (Bernanke et al., 1999; Clarida et al., 1999; Woodford, 2003). Several authors have emphasised the benefits of a nominal anchor within the inflation targeting regime for price stability (for instance, Levin et al. 2004, and Mishkin, 2007). Other authors extend its scope to lower output volatility (Mishkin and Schmidt-Hebbel, 2002, 2007; Fatás et al., 2007). In operational terms, it is assumed that in a country with an inflation targeting regime, expectations must be related to the nominal anchor chosen by the central bank. Hence, the relation between inflation and its expectations must be weak (Levin et al., 2004). Jochmann et al. (2010) have proposed a method to empirically analyse expectations' anchoring. The authors analyse the pass-through coefficient of short- to long-term expectations using breakeven inflation daily data. In a country with anchored expectations, the referred pass-through coefficient must be stable and small; constituting the main result of this investigation. Also, this article makes use of expectations measures focused in more than one horizon. This is worth mentioning since agents could fix their responses to match a longer horizon with the target. This fact is exacerbated with this dataset as its longest horizon coincides with that of monetary policy.

The rest of the article proceeds as follows. Section II reviews recent literature on inflation expectations anchoring, which is scarce for the Chilean case. Section III describes the dataset, comprised by actual inflation and its expectations from the public and from the CBC. Section IV fully describes the econometric analysis and its results. Section V concludes.

II. LITERATURE REVIEW

Several authors have emphasised the role of institutional commitment as a vehicle for price stability through a nominal anchor (for instance, Christiano and Gust, 2000; Mishkin, 2007). In particular, for inflation-targeting countries, setting an explicit figure to which base monetary policy brings attached several areas to be managed by the central bank. These are often referred to as

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1 The original name in Spanish is Encuesta de Expectativas Económicas (literally "Economic Expectations Survey"). An introduction to this survey can be found in Pedersen (2010), and a comparison of its predictive power to other commonly used forecasts is found in Pincheira and Álvarez (2009).

2 These advantages include a long sample, a sufficient number of respondents per month, a rolling event horizon—unlike Consensus Forecasts expectations—, and infrequent changes to the questionnaire. Nevertheless, it still has room for prospective accuracy improvements, as those proposed in Bentancor and Pincheira (2010). A general discussion on the advantages and shortcomings on the use of survey expectations data is provided in Cunningham et al. (2010).

communication and transparency; the building blocks of the credibility needed to succeed at controlling inflation (Geraats, 2010).

Communicational tools go beyond and complement the policy instruments available for policy makers. The former typically refers to the target itself, official economic outlook, parliamentary hearings, open letters, minutes content, and official periodic economic reports. Expectations anchoring refers to the ability of authorities to influence the inflation expectations of the public (King, 2005), which in practice is channelled through a communicational strategy. Anchoring is achieved through both a credibility measure and a latent appraisal of authorities' ability to manage inflation to the target. Hence, firmly anchoring of inflation expectations is equivalent to a "disconnection" between them from actual inflation dynamics, but rather to more sensitivity to central bank communicational developments, considered as policy actions too.

Pierdzioch and Rülke (2013) empirically analyse the role of inflation targeting in 22 countries, including Chile. The authors find more times than not that forecasters scatter their inflation projections away from the inflation target. This result, however, does not necessarily contradict that inflation targets anchor inflation expectations in the long run. It suggests that short-run and medium-run forecasts are prone to, for example, a kind of strategic interaction effects among forecasters. Also including Chile, Gürkaynak et al. (2007) analyse long-run expectations anchoring using event study methodology with daily data–i.e. forward breakeven inflation. Even under this conceptually different econometric setup, the evidence supports that the inflation targeting regime has actually helped anchor long-run expectations.

De Pooter et al. (2014) also analyse long-run expectations anchoring in Chile. Remarkably, the authors make use of both survey- and financial-market-based daily measures of breakeven inflation to find a result like the one provided in this article: anchoring has been more pronounced in the last decade because of improved credibility of the CBC. Also, investors do not seem to systematically alter their inflation expectations because of monetary policy surprises, consumer prices, or activity variables.

Empirical evidence on anchoring is mostly related to successful inflationary experiences under solid policy frameworks. In European countries, van der Cruijsen and Demertzis (2011) examine the aforementioned "disconnection" between actual inflation and expected inflation. The authors find a significant role of the European Central Bank breaking this relationship regarding expectations management. Anderson and Maule (2014) assess the anchoring of short-term inflation expectations in the UK making use of Bank of England forecasts—also a successful anchoring case. Davis (2014) analyses survey data on inflation expectations for 36 developed and emerging countries, finding that in those countries that adopted an inflation targeting regime, the response of inflation expectations to both current and expected inflation or oil price shocks became much less significant and less persistent.

A credible central bank can anchor long-term inflation expectations, but it can also affect short- and medium-term ones through its decisions and communication, especially through inflation forecasts. Pedersen (2015) illustrates the case of the Chilean economy. Empirical estimates suggest that central banks influence the same inflation expectations dataset used in this article, even when controlling for macroeconomic factors. Typically, de-anchoring risks are assessed by focusing solely on long-term expectations, but the responses of long-term inflation to macroeconomic news (for instance, Beechey et al., 2011; Nautz and Stroshal, 2015). Shorter-term inflation expectations have been also studied in the U.S. (Dräger and Lamla, 2013).

Van der Cruijsen and Demertzis (2007) show that the degree of central bank transparency also affects the formation of inflation expectations, while Neuenkirch (2012) and El-Shagi and Jung (2015) extend its scope to short-term interest rate expectations. These findings are relevant because the publication of inflation forecasts constitutes one of the crucial determinants of central bank transparency, as suggested by Eijffinger and Geraats (2006) and Geraats (2010).

Note that managing inflation expectations is crucial for a central bank because of its role in price and wage formation. This fact is emphasised in Paloviita (2008), Forsells and Kenny (2010), and Ciccarelli and Osbat (2017) for European countries, Ball and Mazumder (2015) for the U.S., and Rusticelli et al. (2015) for a set of countries within the *Organisation for Economic Co-operation and Development* (OECD), among many other articles. All these studies make use of a version of the *New Keynesian Phillips Curve* (NKPC) as a baseline framework. For the case of Chile, some reduced-form specification of the *Hybrid NKPC*—including direct measures of expectations in its baseline specification—are found in Medel (2015, 2017) and in Marcel et al. (2017) from a disaggregated services inflation perspective.

All the above-mentioned results rely heavily on how inflation expectations are measured. There is neither a unique nor a widely accepted way to measure the public's inflation expectations. The most common results come from survey-based datasets. However, financial data also provide *direct* measures of expectations. The advantages and shortcomings of both ways are discussed in Cunningham et al. (2010). An *indirect* manner of measuring inflation expectations is to infer the degree of anchoring from a model-based perspective as shown, for instance, in Leigh (2008) for the U.S. economy.

In sum, the anchoring degree depends to a considerable extent on the policy framework. The adoption of an inflation-targeting scheme plays a significant role for expectations anchoring, but so does the transparency and communication strategy conducted by the central bank. These elements are certainly considered in this article, providing an up-to-date assessment of the role of official inflation forecasts and the inflation target in the Chilean economy. Unlike Pedersen (2015), this article also provides an estimation of a model-based implicit-anchored inflation in a unifying simple econometric framework.

III. THE DATA

The dataset is composed of three kinds of inflation data: actual, expected, and forecast. The source of actual inflation is the *National Statistics Institute* (INE) and the CBC for the other series. The sample covers from 2001.9 to 2018.6 (201 observations) in monthly frequency. The start point corresponds to a year after the implementation of a fully-fledged inflation-targeting regime in Chile, and the beginning of the economic expectations survey *Encuesta de Expectativas Económicas* (EEE). Three different consumer baskets are used to conform the current linked-chain series (those of 2008, 2009, and 2013), being 2013 the current base basket.

Inflation expectations are obtained from the EEE. Each month, it reports the median as well as the first and ninth deciles of the distribution of nearly 40 responses (individuals are anonym to the public). For the inflation series, it reports a forecast of the current month, next month, 11-months-ahead, 23-months-ahead, December of next calendar year, and December of two calendar years ahead. A caveat is that, unlike surveys conducted in the U.S. and the Eurozone, the questionnaire does not ask for a horizon greater than that of the monetary policy (24 months). I make use of those expectations at 11- and 23-month-ahead (labelled as *EEE: Infl. 11* and *EEE: Infl. 23*). Forecast errors at both horizons suffer a high level of autocorrelation, and hence, previous forecast errors contain useful information to be used when the current forecast is built.

Official inflation forecasts are obtained from the IPoM. This report is issued quarterly as from after the GFC, and provides inflation forecasts for December of the current year, current-year annual average forecasts, two-year-ahead forecast, and for core inflation at the same horizons. I make use of the forecast for December of the current and next year (labelled as *IPoM: T* and *IPoM: T*+1). The availability of these official forecasts is different from that of the other series, but it contains enough observations (64) to perform non-complex econometric analysis. Pedersen (2015) shows that, as these constitute official forecasts, they influence those of the public (i.e. EEE), exemplifying expectations' anchoring channelling in practice.

Some basic descriptive statistics are shown in table 1.[3] All the time-series are displayed in figure 1. Quite evident is the impact of the commodity price boom and the GFC on the level of the series. Remarkably, this was the period in which more observations of *EEE: Infl. 23* were above 3% (8 observations). The *IPoM: T*+1 forecasts are also close to the target, whereas shorter-term forecast *IPoM: T* follows closely current inflation observations. Most of the analysis considers a split sample between pre- and post-GFC period, whose separating point is the month in which the *Lehman Brothers* bank filed for bankruptcy (2008.9; hence, the pre-GFC sample covers from 2001.9 to 2008.9, and the post-GFC sample covers from 2008.10 to 2018.5).

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3 Note that both kinds of expectations are already released in percentage, while the actual inflation rate is obtained from the Consumer Price Index (CPI), i.e., $Inflation_t = (CPI_t / CPI_{t-12}) \cdot 100 - 100$.

Table 1

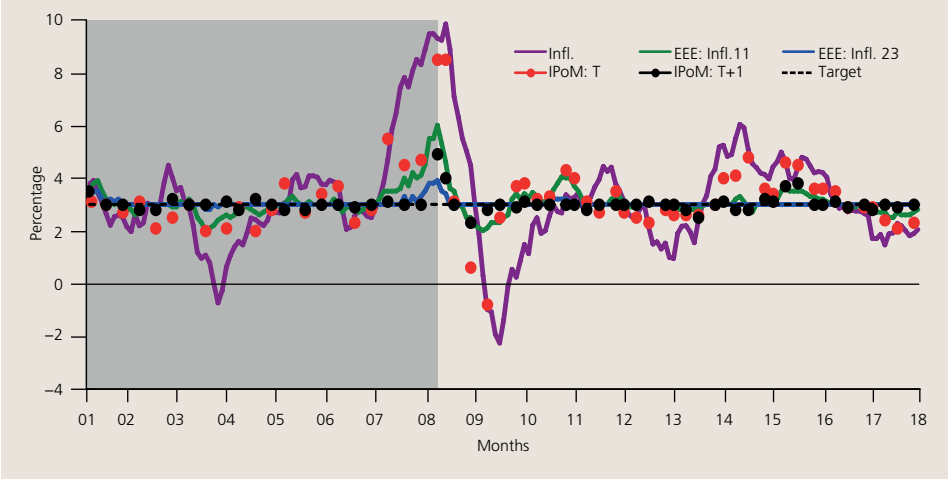
Descriptive statistics of the time series

	Mean	Median	St. dev.	Max.	Min.
	Full sample: 2001.9-2018.5				
Inflation	3.282	3.043	1.985	9.853	-2.266
EEE: Infl. 11	3.083	3.000	0.553	6.000	2.000
EEE: Infl. 23	3.037	3.000	0.137	3.900	2.800
IPoM: Infl. T	3.288	3.100	1.359	8.500	-0.800
IPoM: Infl. T+1	3.064	3.000	0.386	4.900	2.300
	Pre-GFC sample: 2001.9-2008.9				
Inflation	3.445	2.945	2.197	9.476	-0.747
EEE: Infl. 11	3.131	3.000	0.661	6.000	2.000
EEE: Infl. 23	3.066	3.000	0.191	3.900	2.800
IPoM: Infl. T	3.327	2.950	1.467	8.500	2.000
IPoM: Infl. T+1	3.086	3.000	0.437	4.900	2.800
	Post-GFC sample: 2008.10-2018.5				
Inflation	3.141	2.929	2.040	9.853	-2.266
EEE: Infl. 11	3.073	3.000	0.531	6.000	2.000
EEE: Infl. 23	3.023	3.000	0.107	3.900	3.000
IPoM: Infl. T	3.341	3.100	1.585	8.500	-0.800
IPoM: Infl. T+1	3.062	3.000	0.420	4.900	2.300

Source: Author's elaborations based on INE and CBC dataset.

Figure 1

Actual inflation, public expectations, and central bank forecasts (*)



Source: Author's elaboration based on INE an CBC datasets.

(*) Shaded area=pre-GFC sample (20019.9-2008.9). Full sample: 2001.9-2018.5 (201 obs.).

IV. EMPIRICAL ANALYSIS

The same three dimensions of inflation expectations anchoring of Łyziak and Paloviita (2017) are analysed. These are: (i) the response of expectations to current inflation, (ii) the pass-through coefficient from shorter- to longer-term expectations, and (iii) the impact of the inflation target and central bank forecast on expectations. The analysis is then complemented with a time-varying parameter bivariate VAR including current and expected inflation. These same estimates are later used to derive an implicit-anchored inflation series.

Note that the analysis considers the situation in which a central bank operates under an inflation targeting regime, and the inflation target acts as its nominal anchor. It is assumed that the central bank has enough credibility among the public—communicating monetary policy with a high degree of transparency—to interpret deviations of current inflation from the target as merely transitory. Thus, in the long run, the expectations as well as current inflation would return to the target. Besides a high level of transparency regarding future developments of the monetary policy, the central bank could also influence short-term expectations through communicational mechanisms, such as forecasts published in official reports. In this prospective context, de-anchoring effects must be understood as a change in the perception of agents about central bank tolerance to deviations of inflation away from the target. This effect could be translated as an increasing sensitivity of long-term expectations to shorter-term ones, and a higher sensitivity of expectations to current inflation.

In sum, expectations anchoring is a result of policy actions, as inflation fundamentals do not necessarily suggest an underlying prospective inflation always equal to 3%. Therefore, the public (i.e. survey respondents) possibly answer the survey by judging the ability of the central bank to drive inflation towards the target within the policy horizon.

1. Inflation expectations and current inflation

I follow the Ehrmann (2015) procedure by estimating equation (1) in levels of inflation:

$$\pi^{\varepsilon}_{t|t+f} = \alpha + \beta\pi_{t-1} + \varepsilon_t \qquad (1)$$

where $\pi^{\varepsilon}_{t|t+f}$ is the median of responses of EEE expectations at period f made in period t, π_{t-1} is current inflation—lagged one period, corresponding to the latest available observation of current inflation when expectations are made–, and ε_t is an error term.[4] Two versions of this equation are estimated using *EEE: Infl.*

4 *Note that the standard deviation of all estimated coefficients makes use of the Newey and West (1987, 1994) heteroskedasticity and autocorrelation-corrected estimator. In this case, the standard ordinary least squares variance estimator is added to a second term whose size depends on weighted errors' term autocorrelation coefficients. The weights, in turn, are based on an estimated bandwidth used for a (Bertlett) kernel to determine the length of the autocorrelation, under the assumption that it vanishes in time.*

11 and *EEE: Infl. 23*. As mentioned above, if the expectations are anchored, then they must react little and constantly to current inflation. However, and as figure 1 and table 1 report, two different inflation regimes were observed during the sample period. For this reason, I also analyse equation (2):

$$\pi^{\varepsilon}_{t|t+f} = \left(1 - d^{crisis}\right)\left[\alpha^{pre} + \beta^{pre}\pi_{t-1}\right] + d^{crisis}\left[\alpha^{post} + \beta^{post}\pi_{t-1}\right] + \varepsilon_t, \qquad (2)$$

where d^{crisis} is a dummy variable taking the value of 1 from 2008.10 onwards and 0 otherwise. Moreover, the changing nature of the anchoring effect it is commonly accepted in the literature and, precisely by irruptions like the GFC, agents could evolve in their anchoring degree. For this reason, I also depict the β-coefficient in a rolling window scheme for the period 2006.9-2018.5 (i.e. accumulating the first 60 observations available to then move the entire fixed-sized window until the last available observation; finally making use of the 201 available observations).

The estimation results of equations (1) and (2) are presented in table 2. I find that short-term expectations are more sensitive to current inflation than are longer-term expectations, which has a smaller variance. The goodness-of-fit for the *EEE: Infl. 11* equation is nearly double that obtained for *EEE: Infl. 23*. When comparing the size of the coefficients, shorter-term expectations exhibit a coefficient more than five times bigger than that of longer-term expectations (0.183 versus 0.034). This result, however, must be understood statistically as a comparison of the coefficients of two different independent variables, recalling that *EEE: Infl. 23* has by far less variation that *EEE: Infl. 11*. Nevertheless, compared to the results provided by Łyziak and Paloviita (2017) for the Eurozone, these figures could be considered *small* (smallest estimate in Łyziak and Paloviita, 2017, achieves 0.023 with far-long run expectations for the Eurozone).

The GFC appears to have indeed an impact both statistically according to the Wald test and economically according to the estimates. For both types of expectations, the coefficient of lagged inflation decays between 50% and 60% for *EEE: Infl. 11* and *EEE: Infl. 23*, respectively. Remarkably, the inflation coefficient in the regression of each kind of expectations decays after the GFC and, at the same time, the goodness-of-fit coefficient increases, indicating that expectations are indeed better anchored after the GFC.

Table 2

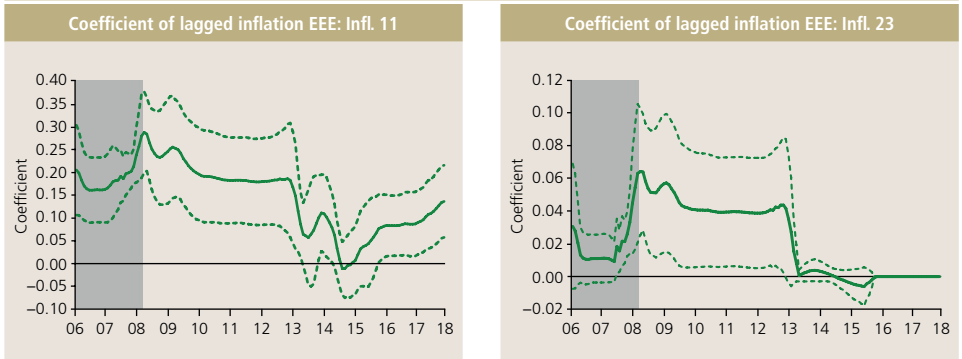
Dependence of inflation expectations on actual inflation (*)

	Full sample (eq. 1)		Full sample with GFC dummy (eq. 2)			Wald test
	β	Adjusted R^2	β^{pre}	β^{post}	Adjusted R^2	NH: $\beta^{pre} = \beta^{post}$
EEE: Infl. 11	0.183	0.467	0.252	0.139	0.505	6.803
	[0.000]		[0.000]	[0.005]		[0.010]
EEE: Infl. 23	0.034	0.253	0.053	0.021	0.317	4.870
	[0.026]		[0.003]	[0.149]		[0.029]

Source: Author's elaboration.
(*) *p*-values in [·] using Newey-West heteroskedasticity and autocorrelation corrected standard errors.

Figure 2

Dependence of expected inflation on lagged inflation, rolling estimates (*)



Source: Author's elaboration.

(*) Rolling window size=60 obs. First estimation sample: 2001.9-2006.8. Shaded area=pre-GFC sample. Confidence intervals=±2 standard deviations.

Aiming directly at analysing the dynamics of expected inflation sensitivity to inflation, figure 2 shows the rolling estimates as a proxy of a time-varying specification. Note that in the Jochmann et al. (2010) sense, a stable low coefficient—far from unity—constitutes evidence of anchored expectations. In particular, figure 2 remarks the impact of the GFC especially when using *EEE: Infl. 23* as the dependent variable. Interestingly, since mid-2013, another anchoring shift is noticed which persists until the end of the sample, plainly fulfilling the definition of inflation expectations anchorage. Finally, the panel II of figure 2 also depicts an indetermination of the estimated coefficient since 2016.5, due to the null variability of the dependent variable (last valid estimation window: 2011.6-2016.5).

2. Short- to long-run expectations pass-through

Again, long-run anchored expectations must not systematically react to shorter-run expectations developments. A stable pass-through coefficient from shorter- to longer-run is also evidence of anchoring, since central banks also use short-term forecasts as a policy instrument (Dräger and Lamla, 2013). To that end, the equation (3) is estimated:

$$\pi^{\varepsilon}_{t|t+f_1} = \alpha + \gamma\pi^{\varepsilon}_{t|t+f_2} + \nu_t, \tag{3}$$

where $f_1 > f_2$, in particular $f_1 = 23$ and $f_2 = 11$, and υ_t is an error term. However, and particularly in this case—where some variation in *EEE: Infl. 23* is noticed—the GFC dummy is also considered (equation 4):

$$\pi^{\varepsilon}_{t|t+f_1} = \left(1 - d^{crisis}\right)\left[\alpha^{pre} + \gamma^{pre}\pi^{\varepsilon}_{t|t+f_2}\right] + d^{crisis}\left[\alpha^{post} + \gamma^{post}\pi^{\varepsilon}_{t|t+f_2}\right] + \nu_t, \tag{4}$$

The results of coefficient estimations are presented in table 3. A pass-through coefficient is obtained near 20% with the full sample, which is also similar to that

obtained with the pre-GFC sample. Remarkably, the coefficient estimated for the period after the GFC drops 63% to 0.155 (supporting a greater anchoring degree); however, not entirely surprising according the dynamics exhibited in figure 1. This period is characterised by low inflation rates—also noticed internationally; see Draghi (2016)—below the target and even near zero. Despite a rapid increment observed from late 2013 until 2014, short-term expectations remained anchored to the inflation target, same as long-term expectations that remained fixed.

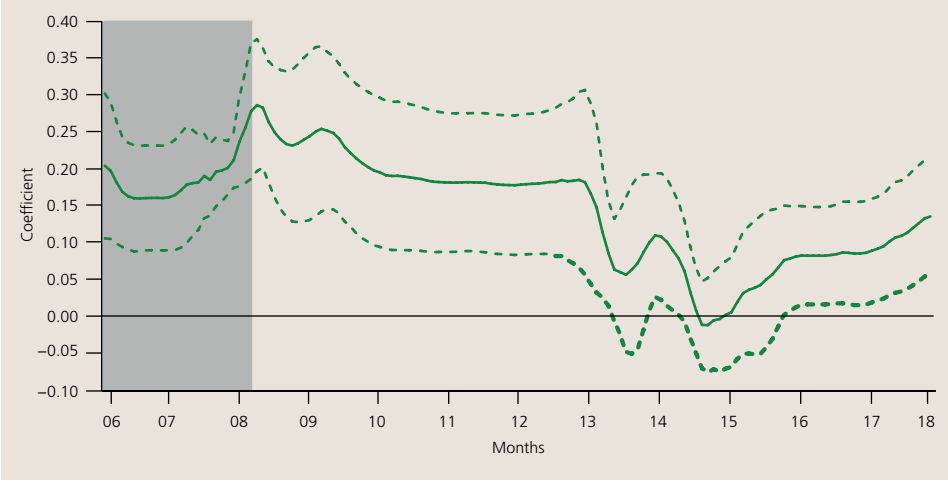
The regressions of table 3 display a relatively high goodness-of-fit coefficient; even in a recursive estimation (not shown; available upon request). This is so because longer-term inflation expectations show very low variation throughout the sample, with some deviation observed during the GFC accompanied by a rise in short-term expectations. As the respondents answer both questions at the same time, the upward revisions coincide between both responses (and not necessarily with current inflation). This allows short-term expectations to have a greater influence than current inflation when explaining long-term expectations. Thus, the regressions in table 3 show higher coefficients and a higher than those in table 2.

Finally, rolling estimates of the γ-coefficient are also analysed in figure 3, for the same period analysed in the previous subsection. Similar dynamics than before is found. In other words, a significant shift due to the GFC followed by a steady estimate surrounding 20%, to then lose statistical significance in 2015, and again become significant but lower than pre-2015 levels. This result is straightforward evidence of inflation expectations anchoring.

Figure 3

Dependence of longer-run on shorter-run expected inflation, rolling estimates (*)

(coefficiente of EEE: Infl. 11 on EEE: Infl. 23)



Source: Author's elaboration.

(*) Rolling window size=60 obs. First estimation sample: 2001.9-2006.8. Shaded area=pre-GFC sample. Confidence intervals=±2 standard deviations.

Table 3

Dependence of longer-run to shorter-run inflation expectations (*)

	Full sample (eq. 3)		Full sample with GFC (eq. 4)			Wald test
	γ	Adjusted R^2	γ^{pre}	γ^{post}	Adjusted R^2	NH: $\gamma^{pre} = \gamma^{post}$
EEE: Infl. 23	0.198	0.640	0.248	0.155	0.683	3.768
	[0.000]		[0.000]	[0.000]		[0.054]

Source: Author's elaboration.
(*) *p*-values in [·] using Newey-West heteroskedasticity and autocorrelation corrected standard errors.

3. Anchoring inflation expectations to the inflation target and central bank forecasts

This subsection analyses inflation expectations anchoring from a slightly different point of view as it includes central bank forecasts as well as the inflation target, both understood as policy actions. Figure 4 depicts the reaction of the public to deviations of inflation from the target and official forecast through a sequence of scatter plots. The way how to read such figures is the following: by taking as an example the upper panels, the Y-axis plots absolute deviations of the expectations with respect to inflation, while the X-axis graphs deviations of inflation with respect to the target. Thus, the 45° line indicates that both deviations would move by the same amount. In other words, expectations accommodate deviations away from the target.

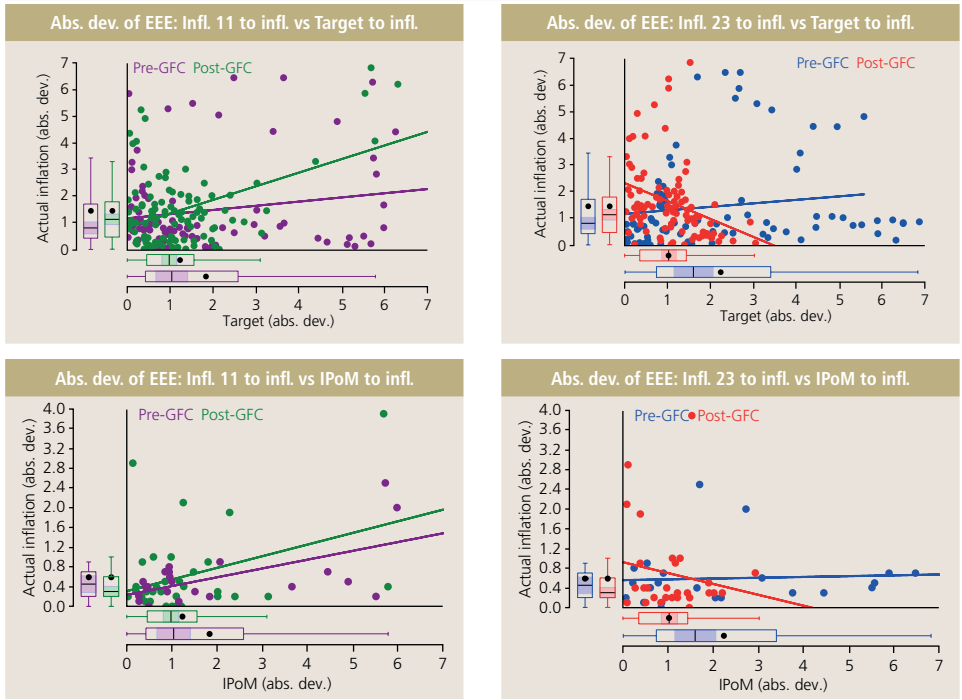
It is important to note that the public (i.e. survey respondents) set their expectations—and, therefore, control the slope of the line—according to their preferences (expectations over or under current inflation). If the line were completely horizontal, it would indicate that the deviations from expectations would be insensitive to the deviations of inflation from the target (perfect anchoring). If the line were to rise above 45°, it would indicate that the expectations would overreact to deviations from the target, suggesting a lower level of anchoring. Finally, a slope less than 45° suggests that expectations react less than deviations from the target, suggesting a higher level of anchoring. A similar logic applies to the case in which deviations are measured with respect to the IPoM (lower panels) instead of the inflation target.

A negative slope is explained by the portion of observations that, during the GFC, reacted late to the rapid change in inflation, which went from 9.9% in 2008.11 to -2.3% in 2009.12, and then to 3.0% in 2011.1. As the IPoM forecasts were more accurate, and inflation approached the target faster, expectations increased their deviation. This, which generates a trend line with a negative slope, is evidence of strong de-anchoring. However, as noted, it is specifically associated with the GFC period.

The results indicate a strong preference for deviations from the target rather than central bank forecasts, particularly after the GFC. In other words, the public reacts more to deviations from the target than to deviations from IPoM forecasts. However, short-term central bank forecasts also proved to be an effective tool for anchoring expectations, similar to the results obtained in Pedersen (2015). Indeed, since *EEE: Infl. 23* has been almost always equal to the target, this also has been an important component to expectations' anchoring effectiveness, as will be discovered next.

Figure 4

Absolute deviations of EEE: Infl. 11 and EEE: Infl. 23 from target versus absolute deviations from current inflation, and absolute deviations of EEE: Infl. 11 and 23 from IPoM versus absolute deviations from current inflation (*)



Source: Author's elaboration.

(*) Pre-GFC sample: 2001.9-2008.9. Post-GFC sample: 2008.10-2018.5.

A more formal analysis is based on the Bomfim and Rudebusch (2000) proposal. Initially, the authors conceive an inflation expectation as a weighted average between the target and lagged inflation, $\pi^{\varepsilon}_{t|t+f} = \lambda^{target}\pi^{target}_{t} + \lambda^{infl}\pi_{t-1}$, but no role is given to central bank forecasts: $\pi^{IPoM}_{t|t+f}$. Hence, and following Łyziak and Paloviita (2017), the analysed equation corresponds to equation (5):

$$\pi^{\varepsilon}_{t|t+f} = \lambda^{target}\pi^{target}_{t} + \lambda^{IPoM}\pi^{IPoM}_{t+f} + \left(1 - \lambda^{target} - \lambda^{IPoM}\right)\pi_{t-1} + \varsigma_t, \tag{5}$$

where π^{IPoM}_{t+f} is the inflation forecast made by the central bank for December of the current year, and ς_t is an error term. Also, note that equation (5) is built to interpret its coefficients as weights, following the original Bomfim and Rudebusch (2000) sense (and in this case, $\lambda^{infl} = 1 - \lambda^{target} - \lambda^{IPoM}$). Equation (5) allows to obtain the degree to which expectations are anchored by computing the figure $\lambda^{target} + \lambda^{IPoM}$. Again, a companion estimation to consider the impact of the GFC is made through equation (6):

$$\pi^{\varepsilon}_{t|t+f} = \left(1 - d^{crisis}\right)\left[\lambda^{target}_{pre}\pi^{target}_{t} + \lambda^{IPoM}_{pre}\pi^{IPoM}_{t+f} + \left(1 - \lambda^{target}_{pre} - \lambda^{IPoM}_{pre}\right)\pi_{t-1}\right] + d^{crisis}\left[\lambda^{target}_{post}\pi^{target}_{t} + \lambda^{IPoM}_{post}\pi^{IPoM}_{t+f} + \left(1 - \lambda^{target}_{post} - \lambda^{IPoM}_{post}\right)\pi_{t-1}\right] + \varsigma_t. \tag{6}$$

Table 4

Impact of actual inflation, inflation target, and central bank forecast on expectations (*)

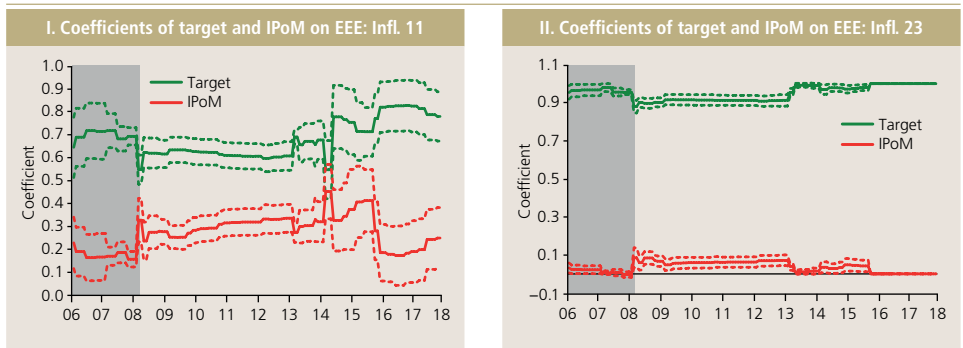
	EEE: Infl. 11 (eqs. 5 and 6)	*EEE: Infl.* 23 (eqs. 5 and 6)
	Full sample: 2001.9-2018.5	
λ^{target}	0.640	0.927
	[0.000]	[0.000]
λ^{IPoM}	0.295	0.055
	[0.000]	[0.028]
λ^{infl}	0.065	0.018
Adjusted R^2	0.744	0.468
	Pre-GFC sample: 2001.9-2008.9	
λ^{target}_{pre}	0.692	0.950
	[0.000]	[0.000]
λ^{IPoM}_{pre}	0.161	0.006
	[0.000]	[0.733]
λ^{infl}_{pre}	0.147	0.044
	Post-GFC sample: 2008.10-2018.5	
λ^{target}_{post}	0.638	0.926
	[0.000]	[0.000]
λ^{IPoM}_{post}	0.324	0.065
	[0.000]	[0.021]
λ^{infl}_{post}	0.038	0.009
Adjusted R^2	0.754	0.483
	Wald test (F-statistic and *p*-value)	
NH: $\lambda^{target}_{pre} = \lambda^{target}_{post}$	0.625	0.538
	[0.433]	[0.466]
NH: $\lambda^{IPoM}_{pre} = \lambda^{IPoM}_{post}$	5.002	3.437
	[0.029]	[0.069]
NH: $\lambda^{infl}_{pre} = \lambda^{infl}_{post}$	8.425	5.076
	[0.005]	[0.028]

Source: Author's elaboration.
(*) *p*-values in [·] using Newey-West heteroskedasticity and autocorrelation corrected standard errors.

Finally, the rolling window estimates of equation (5) are also considered for the same sample used before. The full-sample coefficient estimates are presented in table 4. Remarkably, the biggest weight is assigned to the inflation target—in line with the results of previous subsections—, and this weight in unchangeably dependent on the sample (according to the Wald test). Moreover, the central bank's forecasts play a role especially for short-run expectations (0.295), 46% that of inflation target (0.640). When using long-run expectations, it achieves just a 6% of the weight associated to the target. When the sample is split, weights associated to IPoM and actual inflation are the only ones statistically different, though not economically substantial. What is more important from table 4 is the fact that current (lagged) inflation is not materially important for inflation expectations, neither short- nor long-run.

Figure 5

Weights of inflation target and central bank forecasts on inflation expectations, rolling estimates (*)



Source: Author's elaboration.

(*) Rolling window size=60 obs. First estimation sample: 2001.9-2006.8. Shaded area=pre-GFC sample.

Rolling window estimates shown in figure 5 reveal that the coefficients obtained in table 4 are quite robust. Despite some changes during 2015, the estimates are stable, particularly for *EEE: Infl. 23*. These results reveal the effectiveness of anchoring inflation expectations by the CBC using the communicational instrument of short-term forecasts.

4. Anchors for inflation expectations: empirical estimates

This subsection follows the Demertzis et al. (2008, 2009) approach to derive an empirical inflation target from expectations. The previous approach of Bomfim and Rudebusch (2000) assumes that the inflation target is already an anchor while in this approach the anchor is derived empirically. Consequently, the distance between the estimated and announced targets acts as a central bank credibility check.

The bivariate VAR approach

Demertzis et al. (2008, 2009) assume a bivariate VAR, which contains expected as well as actual inflation, as shown in equation (7):

$$\begin{bmatrix} \pi_t \\ \pi^{\varepsilon}_{t|t+f} \end{bmatrix} = \begin{bmatrix} a_0 \\ c_0 \end{bmatrix} + \begin{bmatrix} a(L) & b(L) \\ c(L) & d(L) \end{bmatrix} \begin{bmatrix} \pi_{t-p_\pi} \\ \pi^{\varepsilon}_{t-p_\pi|t+f} \end{bmatrix} + \begin{bmatrix} \varepsilon_{1,t} \\ \varepsilon_{2,t} \end{bmatrix}, \tag{7}$$

where a_0 and c_0 are intercepts (to be estimated); $a(L)$, $b(L)$, $c(L)$, and $d(L)$ are polynomials containing coefficients of the lagged variables of the VAR (where L is a backshift operator, $L^j x_t = x_{t-j}$), and $\varepsilon_{i,t}$ are white noises. The coefficients of each polynomial , are denoted by $\theta_1, \theta_2, \ldots, \theta_{p\theta}$, where p_θ is the lag length chosen as a block, i.e. $p_a = p_b = p_c = p_d$. In this case, the VAR lag length is chosen as the most frequent suggestion coming from five different criteria: (i) the sequential

modified Lagrange multiplier test, (ii) the final prediction error, (iii) the Akaike Information Criterion (IC), (iv) the Bayesian IC, and (v) the Hannan-Quinn IC. Following the majority rule, the results suggest in two out of five occasions a 7-term lag when using *EEE: Infl. 11*, and in three out of five cases, an 11-term lag for *EEE: Infl. 23*.

The long-run solution of equation (7) delivers equation (8):

$$\pi = \frac{a_0}{1 - a_1 - \ldots - a_{p_a}} + \frac{b_1 + \ldots + b_{p_b}}{1 - a_1 - \ldots - a_{p_a}} \pi^{\varepsilon}, \tag{8}$$

$$\pi^{\varepsilon} = \frac{c_0}{1 - d_1 - \ldots - d_{p_d}} + \frac{c_1 + \ldots + c_{p_c}}{1 - d_1 - \ldots - d_{p_d}} \pi.$$

As stated by Demertzis et al. (2009), the anchoring effect is already based on the notion of Bomfim and Rudebusch (2000) by defining expected inflation as a weighted average between the inflation target (π^*, now a parameter to be inferred) and lagged inflation (becoming $\pi^{\varepsilon}_{t|t+f} = \omega\pi^* + (1-\omega)\pi_{t-1}$). The parameter $0 \leq \omega \leq 1$ measures the degree to which expectations are anchored ($\omega = 1$: perfectly anchored to π^*). By plugging equation (8) in $\pi^{\varepsilon}_{t|t+f} = \omega\pi^* + (1-\omega)\pi_{t-1}$, I obtain the components:

$$\omega\pi^* = \frac{c_0}{1 - d_1 - \ldots - d_{p_d}}, \quad (1-\omega) = \frac{c_1 + \ldots + c_{p_c}}{1 - d_1 - \ldots - d_{p_d}}, \tag{9}$$

and hence, the solution for π^* and the weight in the formation of inflation expectations (ω) are:

$$\pi^* = \frac{c_0}{1 - d_1 - \ldots - d_{p_d} - c_1 - \ldots - c_{p_c}}, \quad \omega = 1 - \frac{c_1 + \ldots + c_{p_c}}{1 - d_1 - \ldots - d_{p_d}}, \tag{10}$$

The estimation results are presented in table 5 for two samples. Note that, given the large number of estimated coefficients (chosen lags are 11 with *EEE: Infl. 11* and 7 with *EEE: Infl. 23*), the full sample estimation is more conservative than that using a shorter sample. Using the full sample, I find an implicit anchor ranging from 3.020 to 3.059 across horizons; virtually equal to the announced target of 3%. When using the post-GFC sample, the same figures are obtained. Interestingly, within this period the maximum weight of expectations formation is achieved. These results emphasise that, despite some variability in expectations, these are coherent with the target when volatile inflation data is factored into the analysis.

The analysis is then complemented with an impulse-response function estimation of current inflation shocks. Following the above-mentioned argumentation, firmly anchored expectations must be invariant to current inflation shocks. Figure 6 shows impulse response functions to +1 standard deviation shocks to current inflation in the VAR specification of equation (7). For the reasons stated above, the preferred results are those estimated with the full sample.

Table 5

Implicit anchors (π^*) and weights in the formation of expectations (ω)

		Full sample	Post-GFC sample
EEE: Infl 11	π^*	3.059	3.067
	ω	0.888	0.985
EEE: Infl 23	π^*	3.020	3.009
	ω	0.969	1.006

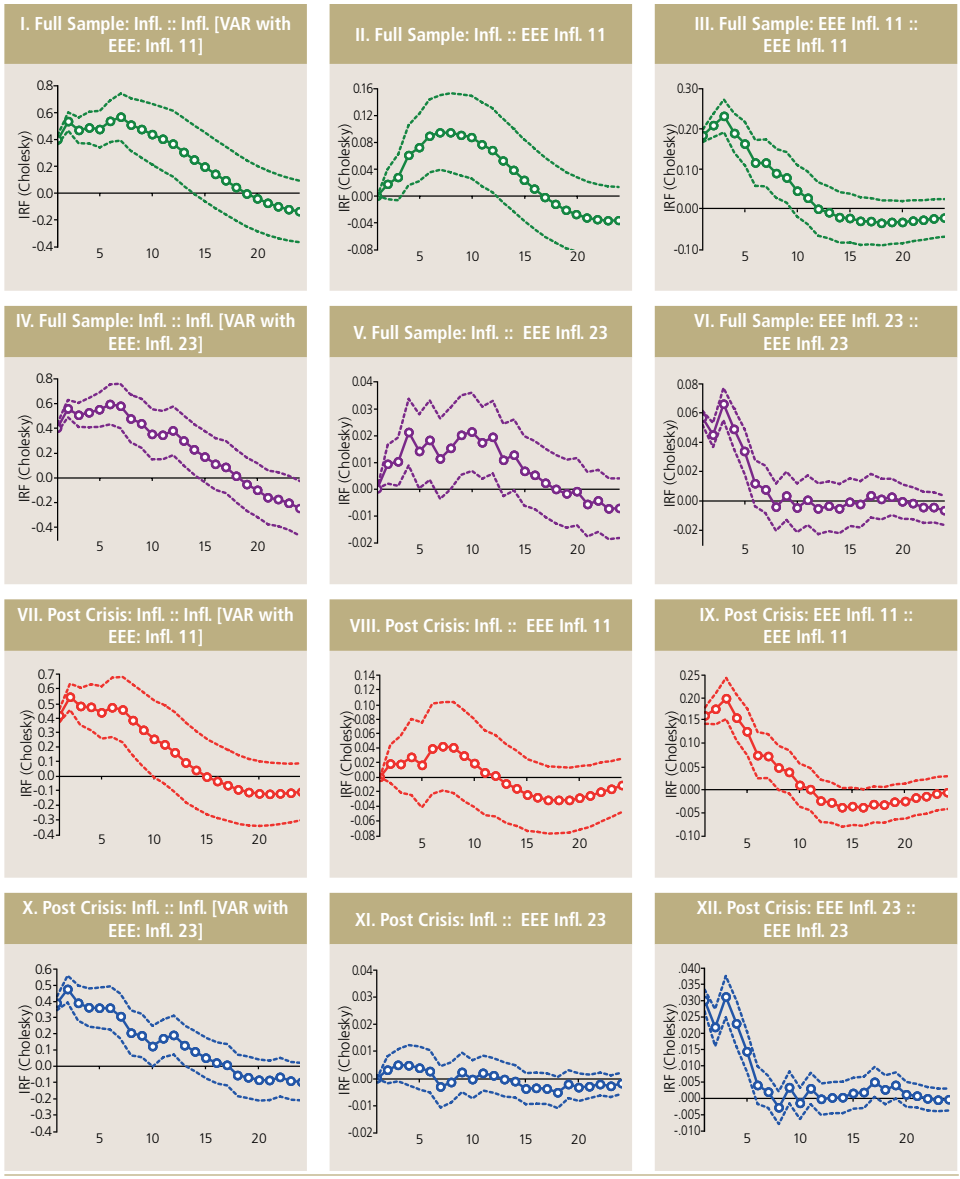
Source: Author's elaboration.

The results of panels I and IV show the response of the inflation shock to itself; representing, thus, a measure of inflation persistence. Panels II and V, in turn, show some statistically significant impact shock from inflation to expectations, peaking at 0.094 and 0.021 basis points, close to 16% and 4% of those related to inflation persistence. However, the biggest impact to expectations comes from the expectations themselves (panels III and VI), peaking 0.233 and 0.066 basis points (41% and 11% of self-inflation shocks). These results, added to those of figure 5, in which changes to the target followed by changes in IPoM forecasts influence the expectations more than headline inflation, leaves a secondary role to recent inflation developments in the formation of expectations. Nevertheless, this also plays a leading role to communicational instruments over current inflation to re-anchor expectations, if needed.

When using the post-GFC sample, no statistically significant pass-through from inflation to expectations is found (panels VIII and XI). This result, while plainly supporting the anchoring of the inflation expectations hypothesis, must be read with caution given the particularly low inflation dynamics in the last part of the sample.

Figure 6

Impulse-response function estimates of expectations on actual inflation (impulse : response) (*)



Source: Author's elaboration.
(*) Full sample: 2001.9-2018.5. Post-GFC sample: 2008.10-2018.5. Bootstrapped confidence intervals=(5%,95%); 5000 replications. Shocks of +1 standard deviation of inflation.

The time-varying parameters VAR approach

Previous VAR-based results are obtained from estimations using the whole sample available without leaving space to rolling estimates as before, supressing the option to discover different monetary policy credibility periods and the implicit-anchor evolution. As these estimates are directly related to coefficients estimation, in this subsection I analyse a time-varying parameter version of the VAR(1) of equation (7), similar to that specified in Demertzis et al. (2008):[5]

$$\begin{bmatrix} \pi_t \\ \pi^{\varepsilon}_{t|t+f} \end{bmatrix} = \begin{bmatrix} a_t \\ d_t \end{bmatrix} + \begin{bmatrix} b_t & c_t \\ e_t & f_t \end{bmatrix} \begin{bmatrix} \pi_{t-p_\pi} \\ \pi^{\varepsilon}_{t-p^{\varepsilon}_{\pi}|t+f} \end{bmatrix} + \begin{bmatrix} \varepsilon_{1,t} \\ \varepsilon_{2,t} \end{bmatrix}. \tag{11}$$

Note that all the time-varying coefficients $\{a_t, b_t, c_t, d_t, e_t, f_t\}$ are estimated using the Kalman filter and are assumed to evolve according to a random walk process. Both error terms are assumed to be independent from each other. Initial parameter calibration is taken from an ordinary least squares estimation using the whole sample.

The results are depicted in figure 7. Overall, the results reveal that estimates are quite stable along the sample, and there are three parameters non-significant: the intercept term (a_t) and the expectations coefficient in the actual inflation equation (c_t,), and that of the lag of actual inflation in the expected inflation equation (e_t,).

Under this scheme, these results mean that persistence is the most critical component of inflation dynamics and expectations. Moreover, the steady estimates of figure 7—and the accommodative swings during 2008—support the point estimates of table 5, suggesting that inflation has been firmly anchored during the analysed sample. In econometric terms, current lagged inflation does not come out as statistically significant due to the collinearity with lagged inflation expectations. However, when excluding the latter, lagged current inflation is as significant as reported in table 2.

A time-varying measure for credibility

In this subsection, previous time-varying coefficient estimates are used to report estimates across the sample of both the implicit-anchor (π_t^*) and the time-varying credibility measure, ω_t.

The results are presented in figure 8. In panel I, jointly with π_t^*, the actual inflation rate and inflation target with ±1% band are also depicted. Remarkably, the green line (π_t^*) deviated from the range in just two episodes (8 observations in total, 4% of the observations): in 2008 jumping to 6.3% and in 2011 barely above 4%. In 21% of the times, the estimated implicit anchor is equal to 3%, and in 80% lies between 2.5% and 3.5%.

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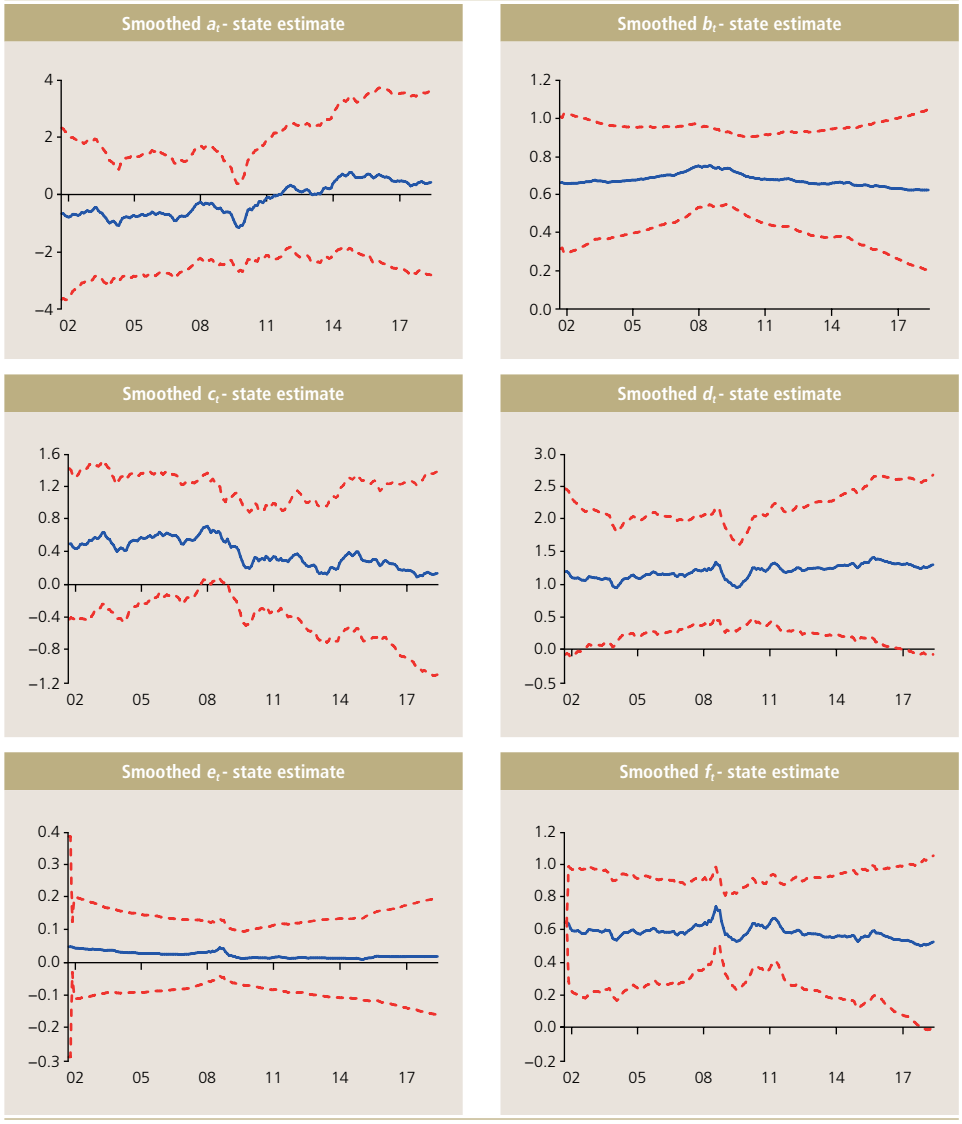
5 Some closer specifications used for the U.S. economy can be found in Stock and Watson (1996), Cogley and Sargent (2005), and Clark and Nakata (2008).

Figure 7

Time-varying parameters estimates of VAR(1) (eq. 11) (*)

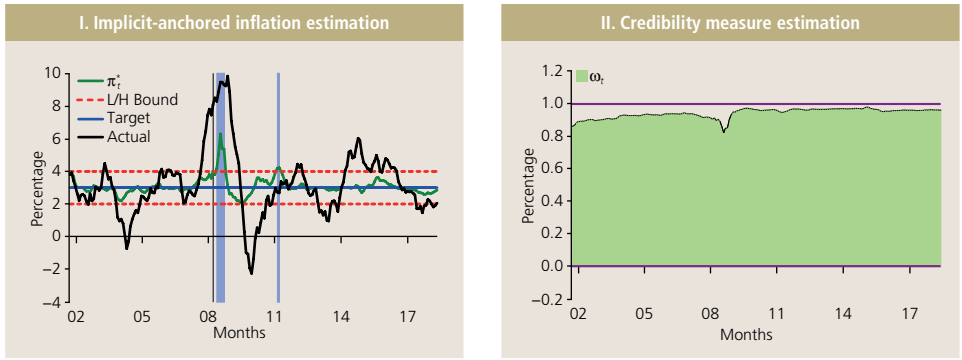


Source: Author's elaboration.
(*) Full sample: 2001.9-2018.5.

It is important to remark that when the implicit anchor deviates from the range it happens by justified common-knowledge reasons. These are the GFC and the rise of commodity prices in 2011; having the common characteristic of being—to some extent—unpredictable shocks.

Figure 8

Time-varying implicit inflation anchor (π_t^*) and credibility estimation (ω_t) (eq. 12) (*)



Source: Author's elaboration.
(*) Shaded areas in panel $I = \pi_t^*$ outside the 3±1% range. Full sample: 2001.9-2018.5.

The attitude of agents towards the CBC's commitment is measured in panel II, as it depicts the ω_t parameter across the sample. This parameter is defined in equation (10) for the VAR(p) model and, in this case, it makes use of time-varying estimates of equation (11), corresponding to:

$$\omega_t = 1 - \frac{e_t}{1 - f_t}. \tag{12}$$

This figure, in line with previous findings, reveal two episodes of inflation anchoring: pre- and post-GFC. From 2001.9 to 2008.9 the average reaches 91.5%, while from 2008.10 to 2018.5 it increases to 96.3% (and 99% of the times above the average of the previous period).

These results are important because they support previous evidence on the existence of two regimes in expectations anchoring and the hysteresis provoked by the GFC on central bank credibility.

V. SUMMARY AND CONCLUDING REMARKS

I performed several econometric testing procedures aiming at analysing to what extent inflation expectations are anchored to the Central Bank of Chile's inflation target. Expectations anchoring is understood as another central bank instrument, in the sense that it helps to promote price stability through the ability of policy makers to influence inflation expectations. Thus, the anchoring of expectations is a result of policy actions.

The econometric analysis relies on the Łyziak and Paloviita (2017) approach, which can be understood as a battery of statistical tests, plus some other robustness exercises found in the related literature. These include the

dependence of expectations on actual inflation, the pass-through of shorter- to longer-term expectations, and the comparison between long-term expectations' deviations and CBC's target deviations. This last feature is extended to consider two communicational tools employed by the CBC, namely the inflation target and forecasts contained in the *Monetary Policy Report*, IPoM for short. Finally, a time-varying parameters VAR model is estimated to derive the implicit inflation target contained in the expectations, plus the anchoring degree to that implicit target. All empirical estimations are based on inflation expectations provided by the EEE survey.

I provide evidence suggesting that agents' expectations are indeed anchored. Estimated pass-through coefficients suggest three anchoring periods. These are (i) the pre-GFC period, (ii) the post-GFC period until mid-2015, and (iii) the low inflation period afterwards.

I find that the pass-through coefficient to short-term expectations from current inflation has declined in time, being null for long-term expectations in the most recent period. The pass-through coefficient from short- to long-run expectations has diminished 65% after the GFC, becoming non-significant in the last period. In a context where most of the times the 23-month horizon expectation is equal to 3%, the public reacts more to deviations from the target rather than to deviations from current inflation.

Estimated implicit-anchored inflation point is found to be equal to the target. Furthermore, time-varying estimates of implicit-anchored expectations lies 96% of the times within the target range of 3±1%. Finally, credibility measures fully support the CBC's commitment with price stability throughout the inflation-targeting period.

Some further research would involve expectations anchoring to a subset of prices, as well as understanding how often and in what manner agents include more information in forming their expectations. These findings could also be compared to those coming from different agents, such as consumers, investors, and domestic versus foreign market participants. Any discrepancy between them will allow to disentangle the effect of communicational tools employed by central banks valuing their underlying effectiveness.

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