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BREAK-EVEN INFLATION RATES: THE ITALIAN CASE

by Alberto Di Iorio[*] and Marco Fanari[*]

Abstract

This paper focuses on break-even inflation rates (BEIRs), a widely used market-based measure of expected inflation, computed from government bonds. In the first part of the paper, we regress the Italian BEIR on several financial variables to assess the contribution of inflation, credit and liquidity components. In the second, in order to disentangle market participants' inflation expectations from risk premia, we estimate a term structure model for the joint pricing of the Italian nominal and real yield curves, considering also credit and liquidity factors. The results show that BEIRs could be a misleading measure of the expected inflation due to the importance of inflation risk premium and credit risk effect. According to our estimates, the decrease in market-based measures of inflation observed in the last part of the sample period seems to reflect a lowering of both inflation expectations and risk premia. Inflation premia co-move with a measure of tail risk of the long-term inflation distribution signalling that investors become more concerned with downside risks.

JEL Classification: C32, E43, G12, H63.
Keywords: inflation-linked bonds, government yields, break-even inflation rate, expected inflation, inflation risk premium, term structure model.
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Contents

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1 Introduction [1]

Government bonds and government inflation-linked bonds (ILB) markets are two of the largest fixed-income markets globally. Over time, ILBs have become available in a number of countries and the euro area has become one of the major sovereign ILB markets, despite a relatively recent start, in terms of both outstanding volumes and turnover. As showed by Table 1, Italian indexed debt is about one third of the euro area's real debt aggregate which amounts to about €540 billion.

Table 1: Nominal and real debt for major euro-area economies as of August 2019.

Country	**Nominal Bond Outstanding**	**Real Bond Outstanding**	**Ratio Real/Nom**
Germany	1,102bn	69 bn	6.3%
France	1,739bn	241bn	13.8%
Italy	1,624bn	168bn	10.3%
euro area	6,900bn	536bn	7.8%

In the euro area, government bonds linked to euro-area inflation are securities that provide investors with protection against consumer price increases. In fact, both the principal to be redeemed at maturity and their coupons are adjusted for inflation in the euro area. As a result of this indexation mechanism, at maturity, the bondholders will recover any loss in purchasing power that occurred during the bond's life. In addition, if there is deflation (i.e. a fall in prices) during the bond's term the amount redeemed at maturity will be no less than the nominal value (the embedded deflation option).

Given their forward-looking nature, asset prices in general and long-term government bond yields in particular incorporate investors' expectations for inflation and future economic activity and risk premia. These premia are the compensation for bearing the uncertainty related to their macroeconomic expectations and should also be expected to vary over time. Long-term nominal bond yields can thus be thought of as comprising three key elements: the expected real interest rate, which is often regarded as being closely linked to expectations for economic activity, the expected long-term rate of inflation and risk premia.

[1]We would like to thank Franco Panfili, Aviram Levy, Stefano Neri, Antonio Rossetti, Tommaso Perez, Dario Ottaviani, Marcello Pericoli, Marco Taboga, Gerardo Palazzo, Simone Letta, Giovanni Secondin, and the participants at the Bank of Italy lunch seminar for their helpful comments and suggestions. The views expressed in this paper are those of the authors and do not necessarily reflect those of the Bank of Italy. All the remaining errors are ours.

ILBs help to complete the market, providing a new asset class whose payoffs are not covered by existing securities. In particular, they offer a true hedge against purchasing power risk. Investors can match ILBs with liabilities that vary with inflation. For example, this is the case for pension funds, insurance companies, and social security schemes, whose liabilities are typically long term and tied to the general price level. In addition, some endowment funds may have a specific mandate to preserve the real value of their capital, and long maturity real bonds represent a natural choice for these long-term investors (Campbell and Viceira, 2001).

Moreover, an ILB allows for the computation of a real yield to maturity and more broadly a real term structure of zero-coupon bond rates across the maturity spectrum. The spread between the nominal and real yields for the same maturity is often referred to as the 'break-even' inflation rate (BEIR), since it would be the hypothetical rate of inflation at which the expected return from the two bonds (nominal vs ILB) are the same. Although BEIRs are usually employed by practitioners as a direct measure of market participants' inflation expectations, using them as an inflation measure can be misleading. BEIRs are risk-neutral inflation measures so they also incorporate inflation risk premia, since investors need to be compensated for inflation uncertainty.

In addition, in a context of high financial integration such as the euro area, with a common currency and monetary policy, it may not be easy to determine whether the inflation expectations embedded in the nominal securities of one sovereign issuer reflect the inflation dynamics of the single country or of the whole area.

As the liquidity of the ILB segment, although improving fast, is likely to remain lower than that of comparable nominal bonds, this may also lead to the presence of a liquidity premium in the yields of ILBs. Moreover, as investors could perceive that ILBs have higher sovereign risk exposure than nominal bonds, BEIRs would also be affected by a credit risk component. Therefore, the presence of liquidity and credit premia could tend to bias the BEIR as a measure of inflation.

Information about expected inflation and the inflation risk premium is valuable for policymakers, investors and sovereign debt managers. In this regard, the presence of a mature market for ILBs is important tool for extracting market participants' inflation expectations. These market-based indicators are complementary to survey-based measures and both are valuable for central banks in assessing their credibility in maintaining stability[2]. Sovereign debt managers (investors) are also interested in the size of the inflation risk premium as they pay (receive) this premium when issuing (buying) nominal bonds. A positive inflation risk premium and the simultaneous absence of any other premium leads to lower costs for the Treasury when it issues ILBs relative to

[2] Corsello et al. (2019) provide an analysis of the use of survey-based measures to investigate the anchoring of inflation expectations to the ECB Governing Council's inflation aim.

nominal bonds (see Garcia and van Rixtel (2007)).

There are various ways of estimating the inflation risk premium on nominal bond yields, thanks to the ample availability of market data. A considerable body of the literature has studied the inflation risk premium using inflation surveys and other measures for inflation expectations. Developments in the research area have focused on also using ILB yields and the associated real interest rate curve. For an overview of this literature see Kupfer (2018) and references therein.

The benefits of using ILB yields are, among others, a high frequency of data and the availability of different maturities. These features are not present in survey inflation forecasts. While this, at first glance, implies the possibility of improved estimates for the inflation risk premium when ILB yields are taken into account, current research shows an added complexity to this apparently obvious result. In fact, some limitations are relevant when the inflation risk premium is estimated using ILBs mainly due to illiquidity and the embedded deflation option. These aspects are considered and taken into account to varying degrees in different studies.

Kupfer (2018) grouped research employing ILB yields to analyse the inflation components into two major categories: regression-based approaches and term structure models. The distinction is made according to whether or not a complete term structure model is specified to estimate the relevant market-based inflation measures. The vast amount of research using ILBs is mainly related to the USA. For the euro area, relevant research on the pricing of the nominal and real bonds is provided by Garcia and Werner (2010), Pericoli and Taboga (2012); Pericoli (2012), Hördahl et al. (2014), Simon (2015), Camba-Mendez and Werner (2017) and Pericoli (2019).

We apply both the above-mentioned analytical tools (regression and term structure models) in order to understand the principal determinants of the dynamics of BEIRs. Starting from an observed persistent and material difference between break-even rates and inflation swap rates, another popular market-based inflation metric, we regress the Italian break-even rate on several financial variables to get some clues about the contribution of inflation components, credit and liquidity factors. In order to disentangle market participants' inflation expectations from their associated risk premia, we estimate a term structure model for the joint pricing of Italian government nominal and real yield curves. This model extends a term structure model developed in the literature for the joint dynamics of the nominal and real yields, also considering a credit factor in addition to the usual nominal and real factors and a liquidity component.

Our main findings are that the Italian break-even rate could be a misleading measure of expected inflation due to the importance of other contributing factors. In our sample covering about ten years (May 2009 - August 2019), inflation expectations are quite stable in the medium to long-term in line with survey forecasts and, since the end of 2013, these expectations have settled at a lower level than the monetary policy target. Inflation risk premia (IRP) have changed sign and take negative values since 2013. Both inflation expectations and risk premia have shown a decline since the second half of 2018 so that the decrease in market-based measures of inflation seems to reflect a lowering of these two components. Moreover, estimated IRP co-moves with a measure of the up vs downside risk of the long-term inflation distribution (balance of risks), signalling that investors become more concerned with downside inflation risks.

Our results are in line with the empirical literature that estimates a down-trending, even negative, inflation risk premium in recent decades. This pattern is consistent with a demand-side shock that pushes inflation and economic activity in the same direction. Under this scenario, a low (or negative) premium could be accepted by investors because nominal assets offer higher real returns. Besides, when the monetary policy is constrained by a binding lower bound, demand shocks may have a larger effect on growth and prices as the central bank cannot respond by adjusting policy rates.

The paper is structured as follows. Section 2 briefly covers the mechanics of the Italian inflation-linked bonds. A regression based analysis of BEIRs is presented in Section 3. The specification and the estimation of the term structure model are presented in Sections 4, while in Section 5 we briefly report theoretical and empirical considerations about the sign of inflation risk premia. Finally in Section 6 we sum up our main findings. Some additional considerations about ILBs and estimated models are provided in the Appendix.

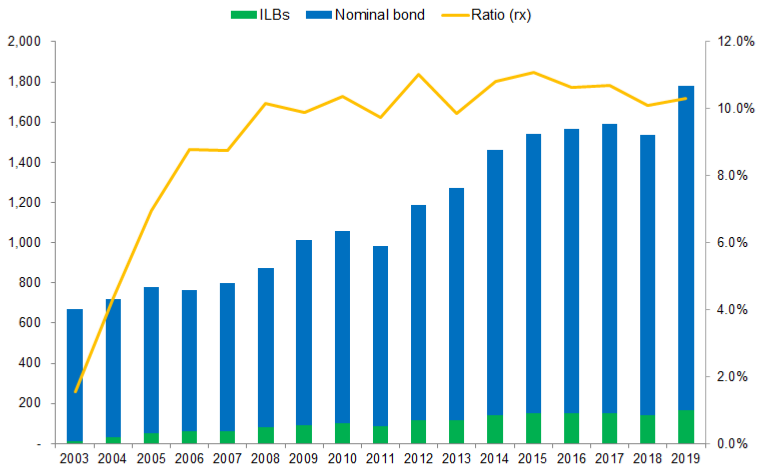
2 The Italian Inflation-Linked Bond Market

Internationally, the role of inflation-linked bonds (ILBs) is growing. Among the major industrialized countries, the United Kingdom was the first to start issuing ILBs in 1981. The US followed in 1997. Countries in the euro area that have introduced ILBs linked to the harmonized European price index are: France in 2001, Greece and Italy in 2003, Germany in 2006 and Spain in 2014. In March 2012, Italy issued the first ILB tied to the national price index (FOI) excluding tobacco ('BTP Italia').

Since the first issuance, the Italian ILB market has been growing almost continuously in absolute terms. Focusing on bonds tied to euro-area inflation, which represents the largest part of Italian indexed bonds, only a few years after the first issuance (see Figure 1), total ILBs outstanding has reached a tenth of nominal issuance. From now on we will refer only to 'BTP€i' (Italian bonds linked to European inflation).

Generally speaking, BTP€i provide constant rates of interest in real terms, that is in terms of purchasing power, fixed at their date of issue (known as the real annual

Figure 1: Evolution of BTP€is outstanding relatively to nominal bonds. Values on the left axis are in million of euros.



coupon rate). The variable amount of the semi-annual coupons is calculated by multiplying half the real annual coupon rate by the nominal principal amount recalculated as at the coupon's payment date. The recalculated nominal principal amount is the subscribed nominal principal amount multiplied by an Indexation Coefficient at the coupon's payment date.

The Indexation Coefficient (IC) is calculated on the basis of the rate of inflation measured by the (unrevised, not seasonally adjusted) Harmonized Index of Consumer Prices for the euro area, excluding tobacco, calculated and published each month by Eurostat. The coefficient allows for the calculation of adjusted values of the nominal principal amount, on the basis of price inflation. Due to the index publication lag, during a generic month m, only the index referring to month $m-2$ or at most to month $m-1$ is available. In order to overcome this limitation, the Reference Price Index is computed using the interpolated value of the index three months and two months prior to the date of interest (coupon payment or principal redemption).

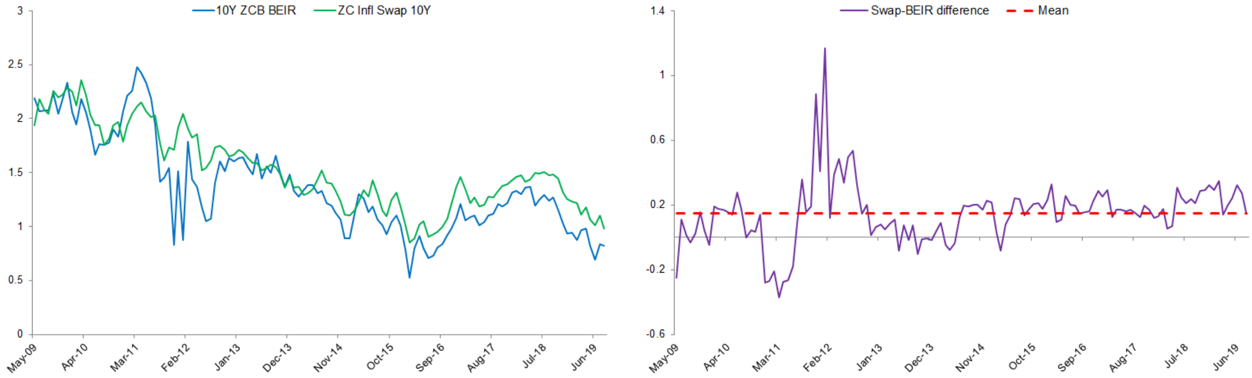
BTP€is have an embedded deflation option, as have many other ILBs; in fact, if the value of the IC for the maturity date is less than one, the amount of the principal redeemed is the nominal value of the bonds. The deflation floor provides protection for the principal but not for the coupon payments. In fact, for these payments, the decrease of the price index from the last coupon date will erode the value of the coupon (compared to the last payment), possibly below the real coupon rate.

The presence on the market of real and nominal government bonds allows the market participants to immediately estimate an expected rate of inflation known as the break-even inflation rate, which is merely the difference between the yield of a nominal and a

Figure 2: Behaviour of market-based inflation measure. Percentage values, monthly data from May 2009 to August 2019.

(a) 10Y Zero coupon break-even rate and 10Y zero coupon inflation swap.

(b) Difference between the two market-based inflation measures considered.



real bond of (approximatively) the same maturity. Another well-known market-based expected inflation measure is given by quotes from zero-coupon inflation swaps, which are derivative contracts where one of the parties pays the other the cumulative inflation over the term of the contract at maturity, in exchange for a predetermined fixed rate. This rate is known as the (zero-coupon) inflation swap rate and it represents a sort of 'synthetic' break-even inflation rate because, if inflation grew at this fixed rate over the life of the contract, the net payment on the contract at maturity would be equal to zero. In the Italian case, the BEIR and the inflation swap rate are rarely in agreement, as shown by Figure 2[3]. Indeed, the difference between these two measures is positive most of the time (14bp in mean) and exhibits its peaks during the euro-area debt crisis. This persistent difference will be investigated in the following sections.

3 Regression-Based Approach

We conduct a preliminary analysis to empirically assess, for the Italian market, the significance and magnitude of factors, affecting BEIRs, mentioned by the ILB literature. The results of this investigation provide us some evidence that Italian break-even rates are not a pure measure of inflation and a guidance in building up in a next step a more structured approach designing an appropriate term structure model.

As mentioned, nominal and real term structures could be used to extract the BEIR for a given maturity. In the absence of liquidity and credit risk premia, the BEIR

[3]To be precise, the BEIR computed in Figure 2 is obtained from zero-coupon bonds so that the nominal and real bonds considered have exactly the same maturity.

represents the inflation compensation requested by investors holding nominal bonds. This compensation is given by expected inflation plus an inflation risk premium. We compute the ten-year BEIR, as the difference between the yield on the ten-year Italian nominal bonds (BTPs) and the yield of the corresponding ILB with the same maturity.

As reported in the literature (see for instance Ejsing et al. (2007), Pericoli (2012)), constant-maturity BEIRs are more accurate than BEIRs computed with yields to maturity. As time passes, the remaining life of a specific nominal yield and a specific ILB yield (nominal and real time to maturities) shortens for a given time horizon (in our case ten years). This shortening is only partially mitigated by data providers substituting old issues with the new bond.

Here we assume that the nominal $y_N^{(n)}$ and real $y_R^{(n)}$ yield of bond of maturity of n, and a zero coupon inflation swap $ZCIS^{(n)}$ with the same maturity can be decomposed as follows:

$$y_N^{(n)} \approx \mathbb{E}\left[r_{t,t+n}\right] + \mathbb{E}\left[Infl_{t,t+n}\right] + \pi'_{cred} + \pi_{Infl} + \pi_{real} \tag{1}$$

$$y_R^{(n)} \approx \mathbb{E}\left[r_{t,t+n}\right] + \pi''_{cred} + \pi_{liq} + \pi_{opt} + \pi_{real} \tag{2}$$

$$ZCIS^{(n)} \approx \mathbb{E}\left[Infl_{t,t+n}\right] + \pi_{Infl} \tag{3}$$

where $\mathbb{E}\left[r_{t,t+n}\right]$ and $\mathbb{E}\left[Infl_{t,t+n}\right]$ are, respectively, the average short-term real rate and inflation rate over the life of the bonds. On the other hand, π'_{cred}, π''_{cred}, π_{liq}, π_{Infl}, π_{opt}, π_{real} are premia for different sources of risk: credit, liquidity, inflation, protection against deflation and the real term premium. It is worth noting that here we are assuming that π'_{cred} and π''_{cred} can be different, that is credit risk can be judged differently for nominal bonds and inflation-linked bonds.

The literature already covered this topic and evidence of different credit risk on nominal and real bonds are found in Goddard and Kitab (2016) and in Simon (2015). In particular, Simon (2015) finds proof of a differing credit risk premium across nominal and real bonds issued from the main euro area countries (especially France and Italy) or, equivalently, a different exposure of nominal and real bonds to sovereign credit risk. Moreover, the magnitude of the credit exposure dominates illiquidity. Here we mention three of the several explanations for this given in the literature. First, Fleckenstein et al. (2014) mentioned a point, often made in the literature on sovereign default risk: the risk in relation to foreign currency-denominated and local currency-denominated sovereign debt may differ. They found this distinction (foreign versus local) is relevant for nominal bonds and ILBs since they mentioned scenarios in which the sovereign issuer might be able to honour its nominal debt by simply 'printing more money', but then not be able to pay off its inflation-linked debt. In essence, they consider ILBs as equivalent to foreign currency-denominated debt from a sovereign default risk perspective. If the market views the default risk of nominal bonds as lower than that of ILBs, then ILBs might trade with higher credit risk premia. Second, Simon (2015) considers the possibility that the harmonized inflation rate gets substantially higher than the issuer's domestic inflation rate. In this scenario, it could become a burden to pay back the indexed bonds tied to the European harmonized index. The third argument is technical, given that real bond cash flows are, generally, skewed towards the maturity if compared with a nominal bond of the same maturity. Indeed, assuming positive inflation, due to the indexation mechanism, the coupon amount paid near the maturity of the bond is larger than that paid at the beginning of the bond's life. In addition, when a bond matures, the bond owner will also receive the principal amount re-evaluated taking inflation into account.

With the previous notation, we can write the BEIR of maturity n as:

$$\begin{aligned}
&BEIR^{(n)} \approx y_N^{(n)} - y_R^{(n)} \\
&\approx \mathbb{E}\left[r_{t,t+n}\right] + \mathbb{E}\left[Infl_{t,t+n}\right] + \pi'_{cred} + \pi_{Infl} + \pi_{real} - \mathbb{E}\left[r_{t,t+n}\right] - \pi''_{cred} - \pi_{liq} - \pi_{opt} - \pi_{real} \\
&\approx \mathbb{E}\left[Infl_{t,t+n}\right] + \pi_{Infl} + \pi'_{cred} - \pi''_{cred} - \pi_{liq} - \pi_{opt} \qquad (4) \\
&\approx ZCIS^{(n)} + \pi_{cred}^{BEIR} - \pi_{liq} - \pi_{opt} \qquad (5)
\end{aligned}$$

where the break-even credit risk premium π_{cred}^{BEIR}, is merely the difference between the nominal and real credit risk premia. Our strategy consists in regressing the 10-year BEIR against some variables that mimic the relevant risk factors shown in (5).

To control for inflation and its associated premium following (4), we use the 10-year zero-coupon Italian inflation swap rate[4]. In fact, this rate reflects both expected inflation over the relevant period as well as an inflation risk premium as assumed in (3).

As a credit risk measure we specify the difference between the nominal 10-year BTP-Bund spread and the 10-year real BTP-Bund spread; we call this variable the cross-country spread (CCS). The rationale of this indicator becomes clear after decomposition (5): the BEIR includes a credit risk component, given by the difference of the credit risk of nominal bonds and the correspondent credit risk of ILBs. If the market perceives the same credit risk, this component is equal to zero, otherwise an additional add-on could become relevant.

We then select a set of variables in order to capture the dimensions of market-wide liquidity and, more specifically, the signs of deterioration in liquidity conditions in the relevant market segments, i.e. nominal bonds and ILBs. Based on the literature, we select two liquidity indicators; while one provides insights into the overall fixed-income market (systematic component), the other reflects specific information about liquidity in the ILB market (idiosyncratic component). The first indicator is the swap-government

[4]The conclusions of a regression analysis would be the same using the European inflation swap rate.

spread (gov), calculated as the difference between the five-ear swap rate and the five-year government bonds rate; Goddard and Kitab (2016) mention it as a Moody's proxy for the illiquidity of fixed-income markets. Besides, as we are interested in the liquidity conditions across the euro-area market, we opt for German data instead of US Treasury rates (as employed by Goddard and Kitab (2016)). They find that a worsening of market liquidity, as measured by the swap-government spread, is associated with an increase in mispricing. The second indicator is the relative bid-ask spread, calculated as the ratio of the bid-ask spread of Italian ILBs to the spread of the nominal bonds. This measure is largely inspired by the liquidity measure specified by Pflueger and Viceira (2011). They use the transaction volume of ILBs relative to nominal bonds for the US and the UK; while we also checked for this related trading activity, considering the volumes associated with all nominal and ILB bonds included in the representative market indices, we found more significant results with the relative bid-ask spread (BA)[5].

As a proxy of the deflation optionality, we consider the price of the quoted 10-year zero coupon inflation floor with a strike equal to zero (ZCF). This inflation derivative is a put option on the inflation rate implied by the price index. This structure pays at maturity the maximum between zero and the difference between the strike price and the observed inflation. In particular, the payoff at maturity depends on the average inflation over the life of the inflation floor over a ten-year period.

The zero coupon inflation floor can thus be used as a hedge against the inflation index ever moving below a certain level during the life of the inflation derivative. Typically, strikes equal to zero or slightly above this threshold have been used during the recent financial crisis to monitor how market participants price the probability of disinflationary and deflationary pressures[6].

In the end, we added a variable to control for market risk aversion, and we used the VSTOXX (Euro Stoxx 50 Volatility), i.e. the volatility index created by the Euro Stoxx and derived from the price of the Euro Stoxx 50 index options for a large spectrum of strikes with a near-term horizon. The VSTOXX ($v2x$) is considered the market's expectation of 30-day forward-looking euro-area equity volatility and is highly sensitive to investors' risk appetite.

[5]The relative bid-ask spread and relative transaction volumes are associated with the constituents of ICE Merrill Lynch Italian government bond indices. Data on individual bonds are aggregated by weighting for their market values.

[6]More broadly, quotes available from inflation floors and caps regarding a spectrum of different strikes (for the same maturity), provide the basis for inferring the whole risk-neutral density of the inflation rate for a given horizon. In other words, this density represents the probability distribution of market participants' inflation expectations.

3.1 Data and Model

In order to implement our strategy, which consists in regressing 10-year zero coupon break-even inflation on some proxy variables as described before, we need the zero coupon nominal and real curves. This is achieved by fitting a Nelson Siegel Svensson model to Italian government bonds using end-of-month values, from June 2009 to January 2019 for a total of 116 observations. All the other variables are available on Bloomberg.

From an econometric point of view, in order to choose the best model to fit our data and for a better understanding the phenomenon of interest, we run an exploratory analysis on our variables. We start with a graphic representation of the time series involved (see Figure 3).

On looking at inflation-related variables, we observe that both inflation swap rates (Figure 3a) and 10-year break-even inflation (Figure 3b) exhibit descending behaviour so that the last sample value of these variables is lower than (swaps) or around (break-even) the level reached during the euro-area sovereign debt crisis. At this point, we limit ourselves to merely describing this behaviour, being unable to attribute this negative trend to a reduction in expected inflation or to a lower inflation risk premium, as the latter is a component of both inflation swaps and inflation break-even as stressed in the previous paragraph.

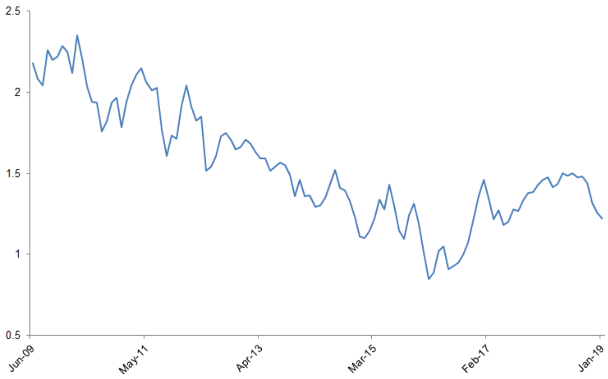
Considering the cross-country spread (Figure 3c), we note that this variable is non-zero, time-varying and generally negative. These considerations provide empirical support for assuming a different sensitivity of nominal and real bonds to credit risk. Starting from the second half of 2010, the spread difference has been steadily negative – with some occasional dates where it turned positive - with the negative peaks reached during the Italian debt crisis, meaning that at a time when Italian credit risk was at its highest level in history, the market put a price on additional credit risk in real bonds than in nominal ones.

Moving to liquidity related variables (Figures 3d and 3e) which, by construction, are positively correlated with a deterioration in liquidity conditions on bond markets, we observe that both showed positive spikes during the euro-area sovereign debt crisis.

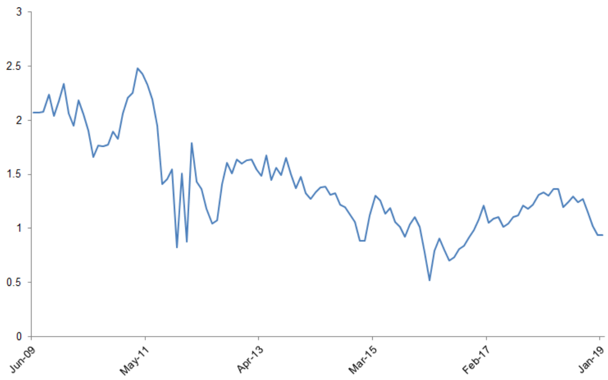
Figure 3f shows, that from the sovereign debt crisis until the end of 2016, deflation was considered a non-negligible event, while nowadays it seems to be less relevant.

Figure 3: Plot of variables considered in regression-based approach. Percentage values, unless BA. Monthly frequency.

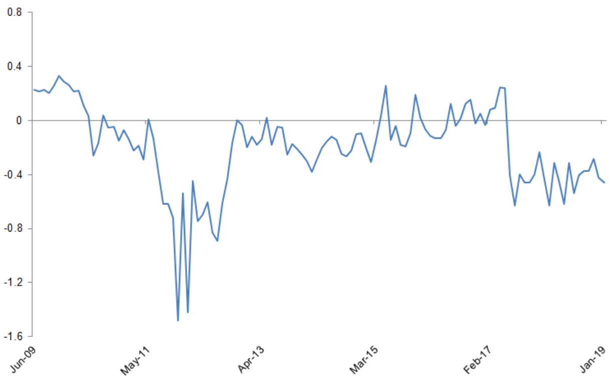
(a) 10-year zero coupon inflation swap ($ZCIS$).



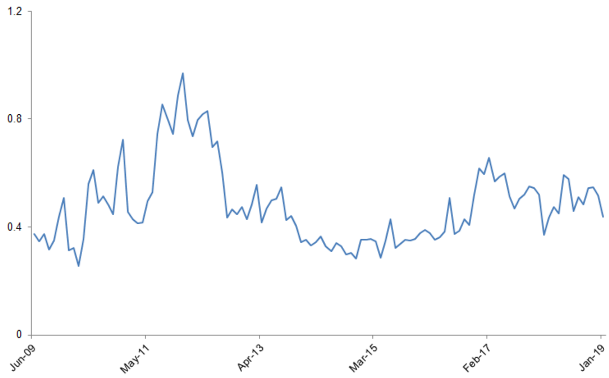
(b) Zero coupon 10-year break-even rate ($BEIR$).



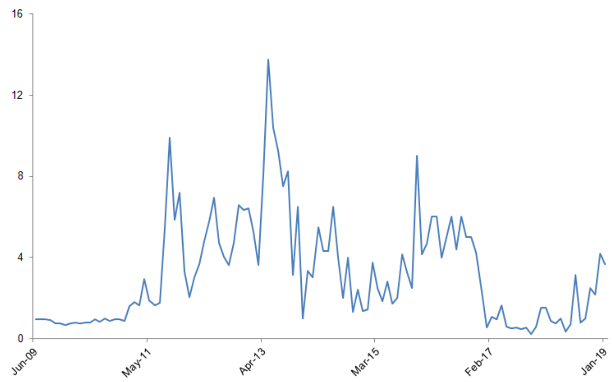
(c) 10-year cross-country credit spread (CCS).



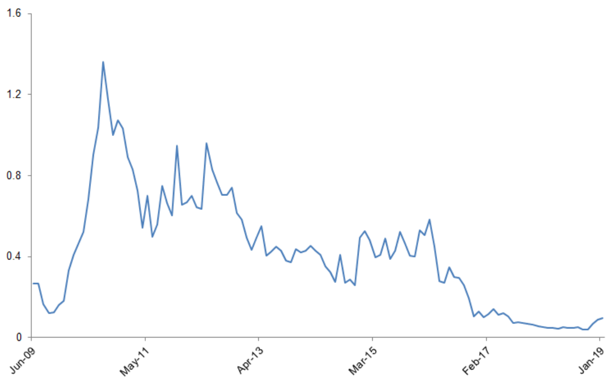
(d) Systemic liquidity measure (gov).



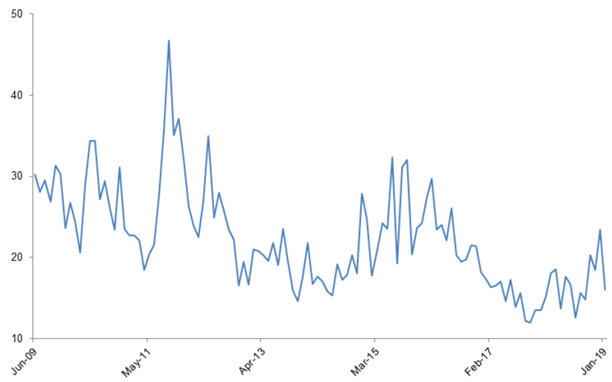
(e) Relative bid-ask spread (BA).



(f) 10-year zero coupon inflation floor 0% (ZCF).



(g) Market risk aversion index VSTOXX ($v2x$).



In order to be relaxed on the order of integration of the time series considered, we use an Auto-Regressive Distributed Lag model (ARDL) with a bounds testing approach as provided by Pesaran et al. (2001). Under mild assumptions (the dependent variable must be $I(1)$ and all the independent variables must have an order of integration of less than 2[7]). This procedure allows us to test if there is a cointegration relationship between variables without focusing too much on what the orders of integration of the variables are[8]. Formally speaking the ARDL model we want to estimate is:

$$\begin{aligned} BEIR_t = \delta BEIR_{t-1} + \alpha_1 CCS_t + \alpha_2 CCS_{t-1} + \alpha_3 v2x_t + \alpha_4 v2x_{t-1} + \alpha_5 BA_t + \alpha_6 BA_{t-1} + \\ \alpha_7 gov_t + \alpha_8 gov_{t-1} + \alpha_9 ZCF_t + \alpha_{10} ZCF_{t-1} + \alpha_{11} ZCIS_t + \alpha_{12} ZCIS_{t-1} + \epsilon_t \end{aligned} \tag{6}$$

With a few algebraic manipulations, it is easy to see that model (6) can be written equivalently as an (unrestricted/generalized) Error Correction Model (ECM):

$$\begin{aligned} \Delta BEIR_t = & \beta_1 \Delta CCS_t + \beta_2 \Delta v2x_t + \beta_3 \Delta BA_t + \\ & + \beta_4 \Delta gov_t + \beta_5 \Delta ZCF_t + \beta_6 \Delta ZCIS_t + \\ & + \theta BEIR_{t-1} + \gamma_1 CCS_{t-1} + \gamma_2 v2x_{t-1} + \gamma_3 BA_{t-1} + \\ & + \gamma_4 gov_{t-1} + \gamma_5 ZCF_{t-1} + \gamma_6 ZCIS_{t-1} + \epsilon_t \end{aligned} \tag{7}$$

To be precise, in the estimated model we have added some dummy variables to control for influential observations [9]. For the sake of brevity we have not inserted them in (7). Following Philips (2018), in order to test cointegration by the Pesaran et al. (2001) bounds test, we have to estimate (7) and assess the absence of serial correlation in its residuals to avoid OLS inconsistency due to endogeneity. All of these steps have been carried out (the diagnostic measures are reported in Appendix B) and the bounds testing approach[10] shows evidence of cointegration at a level of 5 per cent, as can be seen in Table 2.

3.2 Results

The first interesting fact worth noticing in Table 3 is that the coefficient of the lagged dependent variable ($BEIR_{t-1}$) is negative, less than one in absolute terms and with a negligible p-value, i.e. the model is stable and there is a significant equilibrium relationship. In the long run, relationships that are statistically significant involve the cross-country credit spread, market risk aversion, the deflation optionality and the zero coupon inflation swap quote.

[7] We checked these assumptions in the Appendix B; see Tables 4 and 5

[8] Obviously, because we are testing for cointegration, we suspect that variables are $I(1)$.

[9] As a rule of thumb, we classified influential observations such as those in which Cook's distance exceeds $\frac{4}{n-k}$, where n is the number of the observations and k is the number of the parameters to estimate.

[10] The bounds test was performed in R with a *dynamac* package (Jordan and Philips, 2018)

Table 2: Bounds test results.

PSS bounds test results				
	F-test		**t-test**	
Critical value	**I(0)**	**I(1)**	**I(0)**	**I(1)**
10%	1.75	2.87	-1.62	-3.7
5%	2.04	3.24	-1.95	-4.04
1%	2.66	4.05	-2.58	-4.67
	F-statistic = 3.25		t-statistic = -4.44	

Table 3: Error correction model: estimated coefficients.

	Variable	**Coeff.**	**Norm. Coeff.**	**Std. Error**	**t value**	**Pr(>|t|)**	**Significance**
Short run	ΔCCS_t	0.229		0.062	3.676	0.000	***
	$\Delta v2x_t$	-0.003		0.002	-1.267	0.208	
	ΔBA_t	0.000		0.005	0.002	0.998	
	Δgov_t	-0.289		0.132	-2.193	0.031	*
	ΔZCF_t	-0.148		0.119	-1.245	0.216	
	$\Delta ZCIS_t$	0.753		0.109	6.928	0.000	***
Long run	$BEIR_{t-1}$	-0.294	1	0.065	-4.514	0.000	***
	CSS_{t-1}	0.083	0.284	0.040	2.105	0.038	*
	$v2x_{t-1}$	-0.004	-0.014	0.002	-2.229	0.028	*
	BA_{t-1}	-0.001	-0.003	0.003	-0.299	0.766	
	gov_{t-1}	-0.098	-0.335	0.085	-1.158	0.250	
	ZCF_{t-1}	0.106	0.359	0.036	2.921	0.004	**
	$ZCIS_{t-1}$	0.327	1.113	0.078	4.194	0.000	***

The CCS has a positive coefficient, meaning that, other things being equal, the BEIR increases with the widening of the difference between nominal and real BTP-Bund spreads, supporting the empirical evidence of a different credit risk priced by the market, not only across economies but also across fixed-income asset classes. Conversely, a reduction in this spread difference, namely a surge in the real BTP-Bund spread not fully offset by an increase in the nominal one, is associated with a decline in break-even, because the upsurge in real yields is larger than that in nominal yields. The VSTOXX coefficient is negative, suggesting that in the long run, controlling for other variables, an increase in market risk aversion implies a reduction in the break-even rate; this is equivalent to saying that during financial distress (rising market risk aversion), real bonds underperform nominal ones (higher yields). The deflation option only exhibits a significant positive relationship with BEIR in the long-run dynamics, even though the contribution is small; in fact, a permanent surge of 10 bp in the deflation risk proxy implies, on average, a rise of only 3 bp in the break-even rate, controlling for other variables. The positivity of the deflation proxy coefficient is in line with the no-arbitrage arguments: the deflation optionality is an additional protection offered by real bonds pushing its price higher (lowering its yield).

In the end, it is important to highlight that the (normalized) coefficient of the zero coupon inflation swap is slightly above one, suggesting that, by keeping all the other variables fixed, an increase in the swap rate is (more or less) equal to a variation of the same amount for the BEIR. In other words, these two rates seem to move together in the long run when controlling for the effect of the other relevant variables.

In the short run, we have some deviations from the equilibrium relationship, due to the cross-country credit spread and inflation swap rate, but also to the systemic liquidity variable that exhibits a negative relationship with the break-even rate. In other words, when liquidity conditions deteriorate (i.e. our *gov* variable increases), real bond yields are instantaneously pushed higher more than happens for nominal yields.

4 Term Structure Model

The arbitrage-free term structure models specify the (risk-neutral) dynamics of underlying yield curve factors and risk premia under no-arbitrage conditions over time and across bond maturities. The literature started with the models of Vasicek (1977) and Cox et al. (1985).

Affine arbitrage-free term structure models (ATSMs) have become popular with Duffie and Kan (1996): the yields are functions of the underlying factors. While the risk-neutral $\mathbb{Q}$-measure of probability is sufficient for pricing issues, forecasting and decomposing the term structure requires the real $\mathbb{P}$-probability dynamics. A functional form for the market price of risk is required to establish the connection between both measures. Completely affine risk premia (Dai and Singleton, 2000), essentially affine risk premia (Duffee, 2002) or extended affine risk premia (Cheridito et al., 2007) are those most extensively used in the literature.

Abrahams et al. (2015) develop a joint term structure model for nominal and ILB yield curves. Their model is based on the class of essentially affine term structure models and extends the well-known model of Adrian et al. (2013)[11] developed for the nominal yield curve. Their model is quite parsimonious and consists of three nominal pricing factors, two real pricing factors and one liquidity factor. While the three nominal pricing factors originate from a principal component analysis, the two ILB pricing factors are derived via a two-step procedure.

In the first step, real yields are regressed on the principal components for the nominal yield curve. Secondly, the two principal components are derived from the residuals of the first step regression. While an economic interpretation of these factors is not direct, this procedure reduces collinearity. The sixth (liquidity) factor is a composite indicator; specifically, it is an equally weighted average of two liquidity proxies: one is the average absolute yield curve fitting error of US inflation-linked bonds from the Nelson-Siegel-Svensson model, while the other is the ratio of the nominal Treasuries transaction volumes to indexed bond transaction volumes.

The estimation uses excess returns rather than zero coupon yields. Abrahams et al. (2015) assume that return fitting errors are serially uncorrelated; the authors underline that this assumption implies that the only source of predictability in excess returns is captured by the pricing factors, which is a desirable property for a term structure model. Conversely, if the yield pricing errors are i.i.d., then the return pricing errors are serially correlated, signalling that there is a predictability not generated by the pricing factors. The model allows the break-even inflation rate to be decomposed into expected inflation and the inflation risk premium, adjusting for liquidity.

We modify and slightly adapt the research design of Abrahams et al. (2015), considering credit risk as an additional pricing factor, and we apply this specification to the Italian Government's ILBs.

4.1 Data and Model

We obtain nominal zero coupon bond yields from Bloomberg, which are available on daily basis, with tenors $n = 3, 6, 9, 12, 24, \ldots, 120$ months. For estimation purposes, we

[11] The Federal Reserve of New York uses this model as a benchmark for estimating expected average short-term rates and bond term premia. Data are updated daily and are available online (https://www.newyorkfed.org/research/data_indicators/term_premia.html), and also via Bloomberg

need a more granular cross-section of maturities. We obtain the whole spectrum of the relevant tenors, fitting a Nelson-Siegel-Svensson curve (Svensson, 1994) to these data.

The real zero coupon bond yields for Italian ILBs are not available on the common platforms for the required time window. Therefore, the real term structure is estimated by using a Nelson-Siegel-Svensson model on Italian indexed bonds, tied to the HICP and with a residual maturity of more than 24 months. Quotes for these bonds and the Reference Price Index are obtained from Bloomberg. The data run from May 2009 to August 2019.

As shown in the previous regression analysis and according to the literature, liquidity considerations could play a role in the pricing of ILBs. We therefore explicitly model the relative liquidity in the Italian indexed and nominal bond markets. To this end, we construct a liquidity factor, computing the (standardized) ratio of the nominal bond transaction volumes over the ILB transaction volumes; this is one of the constituents of the composite liquidity indicator in Abrahams et al. (2015).

Following our regression-based analysis, we extend the Abrahams et al. (2015) model, considering an ad hoc credit pricing factor, proxied by the BTP-Bund 10-year bond spread.

According to the usual hypothesis of affine models, we assume that pricing factors dynamics under the physical measure ($\mathbb{P}$) follow a VAR process:

$$X_{t+1} - \mu_X = \Phi(X_t - \mu_X) + \nu_{t+1}$$

where ν_t are i.i.d Gaussian with $\mathbb{E}_t[\nu_{t+1}] = 0$ and $\mathbb{V}_t[\nu_{t+1}] = \Sigma$.
The stochastic discount factor, used to price assets, is assumed to be exponentially affine

$$M_{t+1} = exp(-r_t - \frac{1}{2}\lambda_t^T \lambda_t - \lambda_t^T \Sigma^{1/2} \nu_{t+1})$$

where r_t is the nominal short rate and λ_t are prices of risk and both are affine in pricing factors

$$\begin{aligned} r_t &= \delta_0 + \delta_1^T X_t \\ \lambda_t &= \Sigma^{1/2}(\lambda_0 + \lambda_1 X_t) \end{aligned}$$

According to Girsanov's Theorem (Girsanov, 1960), the parameters governing the dynamics of the pricing factors under the risk-neutral measure $\mathbb{Q}$ are defined as:

$$\begin{aligned} \tilde{\mu} &= (I_K - \Phi)\mu_X - \lambda_0 \\ \tilde{\Phi} &= \Phi - \lambda_1 \end{aligned}$$

In ATSMs, the log price, $\log P_t^{(n)}$, of a nominal bond with a maturity n is an affine function of the pricing factors; Abrahams et al. (2015) extend the ordinary ATSMs, assuming that log ILB prices, $\log P_{t,R}^{(n)}$, are also an affine transformation of the pricing factors:

$$\log P_t^{(n)} = A_n + B_n X_t$$
$$\log P_{t,R}^{(n)} = A_{n,R} + B_{n,R} X_t$$

The coefficients $A_n, B_n, A_{n,R}$, and $B_{n,R}$ are defined by recursive relationships obtained by imposing no-arbitrage constraints. This specification implies that the inflation dynamics is linear in the pricing factors:

$$\pi_t = \pi_0 + \pi_1^T X_t$$

Under the pricing measure, the model-implied inflation expectation for a given horizon n is given by the difference between the yields on an n-maturity nominal bond $y_t^{(n)}$ and on a real bond $y_{t,R}^{(n)}$

$$\pi_t^{(n)} = y_t^{(n)} - y_{t,R}^{(n)} = -\frac{1}{n}\left[A_n + B_n X_t - (A_{n,R} + B_{n,R} X_t)\right]$$

The same quantity under the physical measure is obtained by replacing, in coefficients $A_n, B_n, A_{n,R}$, and $B_{n,R}$, the parameters $\tilde{\mu}$ and $\tilde{\Phi}$ with their risk-adjusted counterparts. The difference between these two expected inflation measures constitutes the inflation risk premium.

In our model, we considered eight pricing factors: three factors from the nominal yield curve, three from the real yield curve, one factor to account for liquidity and one that adjusts for credit risk. Following Abrahams et al. (2015), the three nominal factors are simply obtained as the first three principal components of the nominal term structure that are usually interpreted as the level, slope and curvature of the yield curve (see Garbade (1996)). Given that the level of the nominal term structure also depends on the credit risk, in order to avoid overlapping information, the credit factor is obtained as the vector of the residuals of the regression of the BTP-Bund 10-year spread on the three nominal factors. At this point, the real pricing factors are obtained through another orthogonalization process: a regression of the real term structure onto the three nominal factors and the credit factor is conducted. The real factors are then defined as the first three principal components of the residuals of this latter regression. In the end, the matrix X_t is composed of the factors constructed as explained above and of the liquidity factor.

4.2 Estimation

In Abrahams et al. (2015) the VAR process that describes the evolution of pricing factors is estimated by an OLS. It is a known fact that in small samples and in presence of unit roots, the OLS procedure underestimates the true persistence of the autoregressive process. In Bauer et al. (2012) the authors highlight the consequences of the OLS bias when estimating a (dynamic) ATSM. They underline that generally these models may show an OLS bias because of using some pricing factors derived from yields that could be persistent.

In ATSM models, VAR parameters determine the evolution of the risk neutral yields and, when estimated by OLS, these yields converge towards their unconditional mean faster than they would have if the persistence of the model were correctly estimated. These considerations entail that the model implied risk neutral break-even rate is stable so that the biggest part of the variability of the break-even rate is attributed to the IRP. Bauer et al. (2012) review several methods to correct the estimated VAR parameters. In order to evaluate the sensitivity of our results to the persistence bias, we newly estimated our model with corrected (via bootstrap) VAR estimates. The results, that are extensively reported in Appendix C, show a notable variability of the risk neutral break-even inflation, that deviates significantly from the ECB Survey of Professional Forecaster projections (SPF), so that we prefer the OLS estimates. This mismatch between estimates and survey, when correcting OLS estimates, has been underlined also by Wright (2014). He shows implausible short rate dynamics when accounting for bias correction in VAR parameter estimates; this argument is evoked also by Abrahams et al. (2015) where they argue that OLS estimates without correction are fully consistent with survey data.

4.3 Results

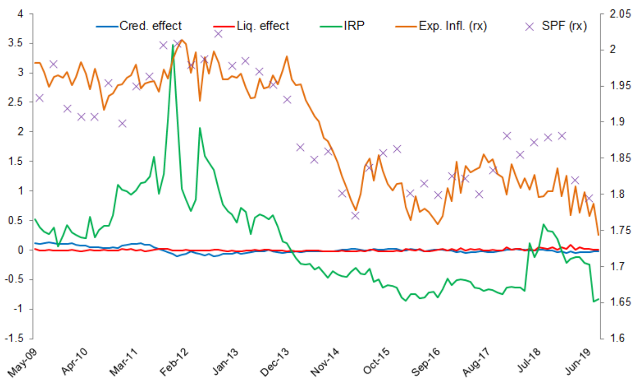
In this work, the ECB Survey of Professional Forecaster expectations is used to provide a proxy of the expected inflation embedded in Italian government bonds. This depends on the limitations of the medium-long term survey data available for Italy. Moreover, although a comparison is made between the estimated inflation expectations and the Survey of Professional Forecasters (SPF) data, an explicit relationship between Italian and euro-area inflation is not assumed. To the best of our knowledge, the only alternative source for medium-long term surveys (for Italy and the area as a whole) is Consensus Economics, which releases data at 6-10-year horizon on a half-yearly basis. Regarding Consensus data, differently from SPF, we do not have the contribution of individual forecasters resulting in a number rounded to the first decimal and in a single value (the central one) instead of the whole inflation distribution.

By comparing the inflation expectations based on the surveys for Italy (Consensus Forecast Italia) and for the euro area (SPF), we detect a substantial alignment of the two series: temporary deviations are generally limited in size and the difference between these expectations is on average less than 1 bp over the sample period.

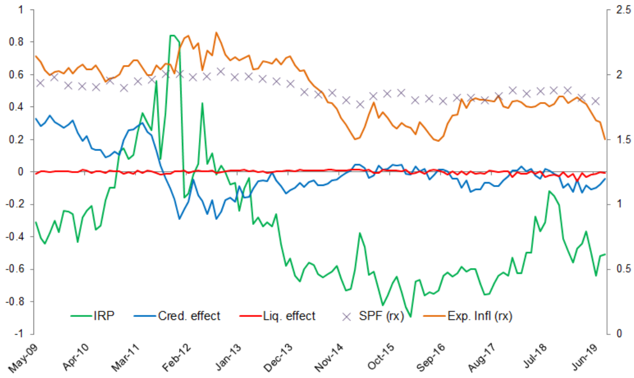
Given the pricing factors and the estimated model parameters, break-even inflation rates at any horizon could be decomposed into expected inflation, the inflation risk premium, the liquidity component, as well as the credit component. We showed the time series of these components for both the 10-year break-even inflation rate as well as the 5-10 year forward break-even inflation rate. The spot 10-year horizon is a canonical maturity in portfolio and risk management, while the medium- to long-term horizon is the usual reference for monetary policy being less influenced by cyclical factors.

Figure 4: Model implied quantities. Percentage values, monthly frequency.

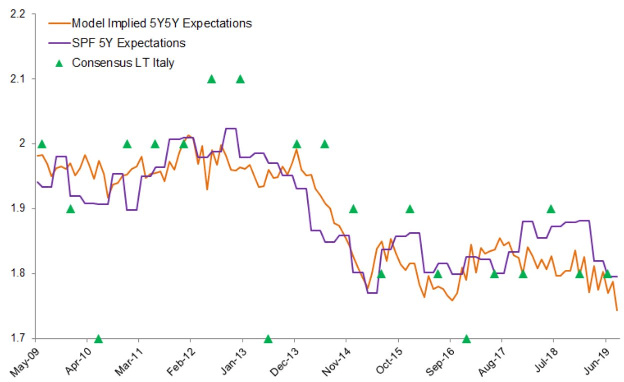
(a) 5Y5Y forward horizon break-even rate decomposition.



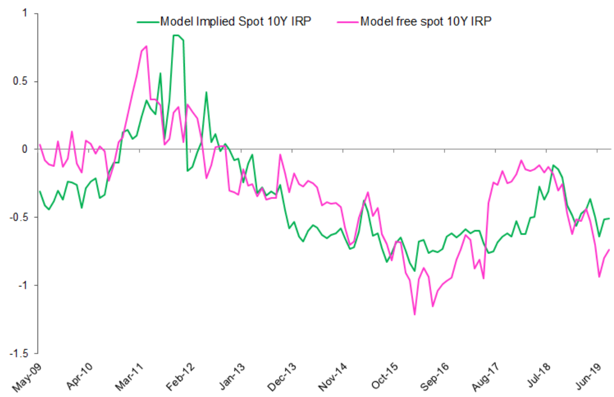
(b) 10Y spot horizon break-even rate decomposition.



(c) Zoomed comparison between 5Y5Y inflation expectations and survey expectations.



(d) Model implied 10Y IRP compared with a 10Y model-free estimated IRP.



Regarding the expected inflation, even though it is quite stable at a medium-to-long horizon and in line with the survey-based expectations, we detect some interesting time variations. Indeed, we find that the 5-10 year expected inflation started to constantly decline from the end of 2013 until the beginning of 2015. Before this downtrend, the estimated inflation was steadily at around 1.95 per cent, while since 2015 it has stabilized at around 1.80 per cent, with further downward pressures in 2019. The same dynamics can be found in the long-term inflation forecast provided by the ECB.

This downward shift is consistent with the results found in Corsello et al. (2019); they prove the statistical significance of two regimes of the mean of the inflation probability distribution obtained by aggregating professional forecasts from ECB surveys. Moreover, the authors find a structural change in the mean in the last quarter of 2013, after which they highlight a lack of a recovery of the expected inflation to the levels prevailing before this break.

Here we want to stress that, as can be seen in Figure 4c, our 5Y5Y model-implied inflation expectation closely tracks the long-term survey projections (Consensus for Italy and SPF for the euro area) although these forecasts are not used in the estimation.

We find a more substantial time variation at a long horizon: the expected inflation associated with 10-year spot rates declines substantially after 2013, reaching minimum levels of around 1.5 per cent in the following three years. Only in the second half of 2016 did the expected inflation tend to stabilize around at 1.7 per cent (see Figure 4b).

Figures 4a and 4b show that the 5Y5Y-forward compensation for bearing inflation risk (i.e. the estimated inflation risk premium, adjusted for liquidity and credit components) is positive or essentially zero for the first years (up to 2013). It then becomes negative for all the remaining periods, with a sudden rise at the end of the first half of 2018, largely offset in the following months. The 10-year inflation risk premium exhibits the same dynamics of the medium-to-long term one with some differences in the levels.

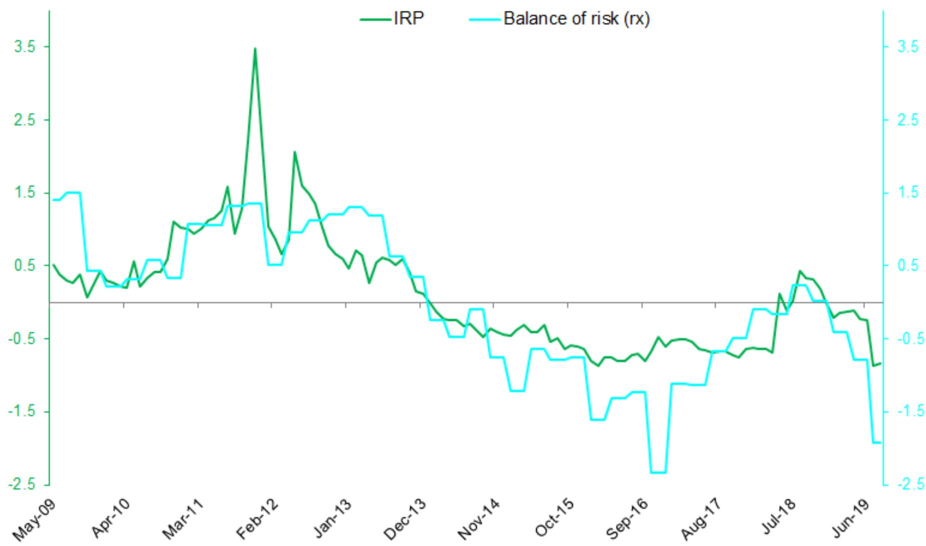
In addition, our estimate of the 10-year spot inflation risk premium is in line with a model-free estimate (see Figure 4d), computed as the difference between the zero coupon break-even rate on the one hand and the inflation expectation based on the European Central Bank Survey of Professional Forecasters (SPF) and the credit spread on the other hand.

The credit component represents the third contributor to the break-even rate (see Figures 4a and 4b), supporting the empirical evidence of the relevance of this factor in the pricing of ILBs vis-à-vis nominal bonds. Even though this relative pricing factor stands at a low level, its dynamics over time are informative. In fact, its positive values taken in the first years of our sample imply a relative cheapness of the nominal bonds compared with the real ones. This dynamics is different in the following years: the negative sign of this credit risk component has been associated with stressed markets conditions for Italian sovereign bonds.

The liquidity effect seems have played only a minor role: the contribution of this component over time is very small.

In order to interpret the estimated inflation risk premium, inspired by Camba-Mendez and Werner (2017), we construct an indicator of up vs downside tail risk for the long-term inflation distribution derived from the SPF[12]. Specifically, our balance of risks is defined as the (standardized) difference between the probability of inflation being higher than 3 per cent and the inflation distribution being under the 1 per cent threshold[13]. We found a strong positive co-movement between the medium- to long-term IRP and the inflation balance of risks. As can be observed in Figure 5, the balance of risks around the inflation outlook moved to the downside around the end of 2013 and stayed there until the end of the sample period, meaning that in this period, investors have become more concerned with downside inflation risks. This is consistent with the fall of the estimated inflation premium into negative territory in the same period. It is worth noticing that this measure of tail risks is approaching the minimum reached in 2016.

Figure 5: Balance of risks and 5Y5Y estimated IRP. Left axis values are in percentage, data are at a monthly frequency, balance of risks data are constant over each quarter.



[12]We obtained an aggregated distribution of the long-term inflation forecast by averaging the distributions given by each forecaster.

[13]Camba-Mendez and Werner (2017) define the balance of inflation risk as the difference between the price of a one-year-ahead inflation cap with a strike price of 4 per cent and a one-year-ahead inflation floor with a strike price of 0 per cent.

5 Interpreting Inflation Risk Premia

The literature is currently focusing on a possible change in the sign of the inflation risk premium. Theoretically, the sign depends on the correlation with the marginal intertemporal rate of substitution for the consumption of the representative investor. We will briefly ouline this argument; for an overview of this point, see Cochrane (2011), Kupfer (2018), Christensen et al. (2010) and the references therein.

In an arbitrage-free setting (ensuring consistent pricing), the price of a nominal bond at time t that pays one dollar at time T (payoff in nominal terms) $P^N_{t,T}$ is the (risk neutral) expectation of the ratio of the nominal stochastic discount factor at time $t+T$ (M^N_{t+T}) over the same discount factor at time t (M^N_t). Similarly, the price of a ILB (real bond) that pays one unit of the consumption basket at time T (payoff in real terms) $P^R_{t,T}$ is the ratio of the real stochastic discount factors for the same maturities: $t+T$ (M^R_{t+T}) and t (M^R_t).

Formally:

$$P^N_{t,T} = \mathbb{E}_t\left[\frac{M^N_{t+T}}{M^N_t}\right], \qquad P^R_{t,T} = \mathbb{E}_t\left[\frac{M^R_{t+T}}{M^R_t}\right]$$

For pricing consistency, the nominal stochastic discount factor for each maturity is given by the real discount factor divided by the price level Q_t for the same time horizon, so that it holds that:

$$\begin{aligned}
P^N_{t,T} &= \mathbb{E}_t\left[\frac{M^N_{t+T}}{M^N_t}\right] = \mathbb{E}_t\left[\frac{M^R_{t+T}}{M^R_t}\frac{Q_t}{Q_{t+T}}\right] = \\
&= \mathbb{E}_t\left[\frac{M^R_{t+T}}{M^R_t}\right] \times \mathbb{E}_t\left[\frac{Q_t}{Q_{t+T}}\right] + \mathbb{COV}_t\left[\frac{M^R_{t+T}}{M^R_t}, \frac{Q_t}{Q_{t+T}}\right] = \\
&= \mathbb{E}_t\left[\frac{M^R_{t+T}}{M^R_t}\right] \times \mathbb{E}_t\left[\frac{Q_t}{Q_{t+T}}\right] \times \left(1 + \frac{\mathbb{COV}_t\left[\frac{M^R_{t+T}}{M^R_t}, \frac{Q_t}{Q_{t+T}}\right]}{\mathbb{E}_t\left[\frac{M^R_{t+T}}{M^R_t}\right] \times \mathbb{E}_t\left[\frac{Q_t}{Q_{t+T}}\right]}\right).
\end{aligned}$$

Moving to yields:

$$\begin{aligned}
y^N_{t,T} &= -\frac{1}{T}\ln P^N_t \\
&= -\frac{1}{T}\ln \mathbb{E}_t\left[\frac{M^R_{t+T}}{M^R_t}\right] - \frac{1}{T}\ln \mathbb{E}_t\left[\frac{Q_t}{Q_{t+T}}\right] - \frac{1}{T}\ln\left(1 + \frac{\mathbb{COV}_t\left[\frac{M^R_{t+T}}{M^R_t}, \frac{Q_t}{Q_{t+T}}\right]}{\mathbb{E}_t\left[\frac{M^R_{t+T}}{M^R_t}\right] \times \mathbb{E}_t\left[\frac{Q_t}{Q_{t+T}}\right]}\right) \qquad (8) \\
&= y^R_t + \pi^{exp}_t + IRP_t \qquad (9)
\end{aligned}$$

The decomposition of the nominal bond yield y^N in (9) shows that it consists of three terms: (i) the real yield y^R, (ii) the expected inflation π^{exp} and (iii) the inflation risk premium IRP_t.

From (9), it is clear that the sign and the magnitude of the inflation risk premium depend on the covariance between the real discount factor and the purchasing power (the inverse of inflation). The IRP will be positive if and only if this covariance is negative.

$$IRP > 0 \Longleftrightarrow \mathbb{COV}_t \left[\frac{M_{t+T}^R}{M_t^R}, \frac{Q_t}{Q_{t+T}} \right] < 0$$

In other words, the IRP is positive if and only if the real discount factor tends to be high at the same time that price inflation is high. That is to say, the riskiness of nominal bonds depends on the covariance between real stochastic discount factors and inflation.

As observed by Camba-Mendez and Werner (2017), in a consumption-based asset pricing model, discount factors are linked to the rate at which the investor is willing to substitute consumption today for consumption tomorrow. This marginal rate of substitution depends on the future marginal utility of consumption. Assuming that the utility function of consumption is concave (i.e. the marginal utility is higher when the level of consumption is low), there is a positive premium for investing in nominal assets when the future inflation is expected to be positively correlated with the future marginal utility of consumption. In other words, a financial asset that provides returns eroded by high inflation when more wealth is desired should only be held if it offers a positive premium. On the other hand, when future inflation is expected to co-move negatively with future marginal utility, a negative premium could be accepted because nominal assets offer higher real returns.

Going further, a positive correlation between consumption growth and inflation (negative IRP) occurs when the economy is expected to be hit by a shock on the demand-side that pushes inflation and economic activity in the same direction. A typical scenario could be a recession largely attributable to a weak aggregated demand that keeps inflationary pressures subdued. Vice versa, when shocks affect supply, the correlation between consumption and inflation becomes negative (positive IRP) that is consistent, for instance, with a stagflationary macroeconomic environment.

Moving to empirical territory, Chernov and Mueller (2012) show that the inflation risk premium in the United States has trended downwards since the mid-1980s, by comparing different model specifications. D'Amico et al. (2018) also estimate a down-trending inflation risk premium that turns negative around financial crises, similarly to findings in Grishchenko and Huang (2013). Another important contribution is due to

Chen et al. (2016) who, in order to investigate the sign of the inflation risk premium, estimate the correlation between forward consumption growth and long-run inflation. They show that this correlation has not been stable over time in the US. It was deeply negative in the 1980s, consistently with the supply-side shocks that raised inflation and weakened economic growth, suggesting a positive IRP at that time. The correlation then trended up over time and switched sign with the financial crisis, meaning that the IRP turned negative around this period, consistent with interpreting the financial crisis as a large negative demand shock as noted by Ferrero and Neri (2017). Macroeconomic theory suggests that when the zero lower bound (ZLB) does not bind, the central bank can offset demand fluctuations by acting on the interest rate level, which means that demand shocks may have little effect on inflation or economic activity. Vice versa, when monetary policy is constrained by a binding lower bound, the central bank cannot respond by adjusting policy rates so that demand shocks may have a greater effect on growth and prices. To evaluate the effect of the ZLB on the inflation risk premium, Gourio and Ngo (2016) develop a fairly standard New Keynesian macroeconomic model that generates positive term and inflation risk premia when policy rates are far from the lower bound, while these premia fall when the economy is close to the ZLB.

6 Conclusions

This paper aims to understand the principal determinants of the Italian break-even inflation rate. To this end, we used two of the main approaches suggested by the literature on this subject. Specifically we set up an ARDL model and regress break-even rate on a set of variables that proxy inflation, market risk aversion, protection against deflation, credit as well as liquidity risk to get some insights into the importance of these factors. Subsequently, in order to disentangle market participants' inflation expectations from their associated risk premia, we estimate a term structure model for the joint pricing of the Italian government's nominal and real yield curves. This model extends a well-known no-arbitrage Gaussian affine specification also considering a credit pricing factor, in addition to the usual nominal and real factors and a liquidity component.

We find that the Italian break-even inflation rate could be a biased measure of the expected inflation due to the importance of other contributing factors mainly given by inflation risk premia and credit effect. In our results, the expected inflation is quite stable at the medium to long-term horizon and in line with the survey expectations, but exhibits some time variations. In particular, inflation pattern shows two regimes: the first one, before 2013, in which expectations move around the ECB definition of price stability and the second one, which started in 2015, in which inflation settles at a level below the ECB Governing Council's inflation aim. Furthermore, since the second half of 2018, inflation shows a further decline in line with the (in-sample) historical minimum reached in 2016. The estimated inflation risk premium changes sign and becomes steadily negative since the end 2013. In addition, in order to test the reliability of our estimate, we compare our model-based IRP with a model-free indicator obtained by subtracting the expected value of the SPF inflation forecasts from the break-even inflation rate, adjusted for credit risk. We find these two measures co-move strongly signalling that our model provides reasonable estimates. Similarly to expected inflation, our estimated IRP also shows a decline in the last part of the sample period, so that the decrease in market-based measures seems to be attributable to a lowering of both the premium and inflation. Finally, further support for our findings is given by the high correlation between the model-implied IRP and the inflation balance of risks. The latter indicator, which is a measure of up vs downside tail risk, is close to its minimum value, warns that investors have become more concerned with downside inflation risks.

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Appendix A

The rationale of inflation-linked bonds

Historically, the issuance of ILBs responds to a number of economic and financial arguments that could have positive or negative implications in terms of social welfare, debt management, the implementation of monetary policy, and capital markets. These arguments are highly complex and we have only explored a selected set here drawing on the detailed reviews provided by Price (1997) e Garcia and van Rixtel (2007).

First, in a high and volatile inflation environment, ILBs were the best available option for raising long-term funds in financial markets. Emerging markets (like Chile, Brazil, Colombia and Argentina) and European countries (like France and Finland) issued ILBs for this reason. In 1983, Italy raised capital with a ten-year maturity, at a time when it was experiencing some difficulties in issuing nominal bonds with long maturities.

Second, ILBs provide a device to those countries who want to signal a strong commitment to low inflation and in that case, the issuance of ILBs reduces the incentive to ease the real value of sovereign debt via higher inflation. Along these lines, Margaret Thatcher argued that inflation-linked debt was a 'sleeping policeman' that helped to control inflation.

Third, the Treasury could reduce its funding costs by not paying the inflation risk premium, provided that the inflation risk premium is positive, and/or capitalizing on the excessive inflation expectations embedded in the nominal bonds rates. In the context of disinflationary policies, a group of countries (like the United Kingdom, Australia, Sweden and New Zealand) issued indexed debt to add credibility to their governments' commitment to these policies and to reduce their borrowing costs.

Fourth, inflation-linked bonds help to complete the market, providing a new asset class whose payoffs are not covered by existing securities. In particular, they offer a true hedge against purchasing power risk. Investors can match ILBs against liabilities that vary with inflation. Examples of indexed debt are provided by pension funds, insurance companies and social security schemes, whose liabilities are typically long-term and tied to the general price level. In addition, some endowment funds may have a specific mandate to preserve the real value of their capital and the long maturity real bonds represent a natural choice for these long-term investors (Campbell and Viceira, 2001).

Fifth, an argument on moral grounds that would justify the existence of ILBs was expressed by Milton Friedman. In his view, 'The government cum monetary authority created inflation in the first place and therefore has the responsibility to provide means by which citizens can protect their wealth'. By indexing its liabilities, the government would give back to the public part of the inflation tax for which it was responsible. From another point of view, the Friedman motivation can be seen as a distributional argument; in fact, unanticipated inflation implies the unintended transfer of wealth from lenders to borrowers, and indexed bonds protect lenders from this effect.

Sixth, ILBs provide policy makers, as well as investors, with new tools for gauging the public's inflation expectations or more precisely the expectations of market participants. This kind of information, extracted using ILBs, needs to be carefully considered. Indeed, this risk-neutral market-based measure also includes risk premia that have to be disentangled from pure inflation expectations. In any case, risk premia also provide valuable information to policy makers as well as to investors.

Appendix B

Regression approach

B.1 Preliminary analysis

In this section, we report the results of stationarity tests on the variables considered in the regression model (6). We performed both a Dickey-Fuller test (Dickey and Fuller, 1979) - whose null hypothesis is non-stationarity - and a Kwiatkowsky, Phillips, Schmidt and Shin (KPSS) test that tests stationarity[14]. We tested the variables in terms of both levels and differences in order to check the ARDL assumptions (dependent variables of order $I(1)$ and independent variables of order of less than 2). The $\alpha = 5\%$ and $\alpha = 1\%$ confidence levels were taken into account.

The results are reported in Tables 4 and 5. Looking at the break-even rate, which is the dependent variable, we see that both tests conclude the non-stationarity of the level series and the stationary behaviour of the difference series, i.e. the BEIR is an $I(1)$ variable. For all of the independent variables, tests on difference series agree on the stationarity, supporting the assumption that the explicative variables have an order of integration lower than 2.

[14]All the tests have been conducted with the *urca* package in R (Pfaff, 2008).

Table 4: ADF test on model variables.

Variable	**Levels Test. stat.**	**Differences Test stat.**
$BEIR$	$\tau_3 = -2.55$ $\Phi_2 = 2.44$ $\Phi_3 = 3.30$	$\tau_3 = -7.82$ $\Phi_2 = 20.38$ $\Phi_3 = 30.57$
CCS	$\tau_3 = -2.60$ $\Phi_2 = 2.35$ $\Phi_3 = 3.40$	$\tau_3 = -9.12$ $\Phi_2 = 27.75$ $\Phi_3 = 41.63$
$v2x$	$\tau_3 = -3.97$ $\Phi_2 = 5.29$ $\Phi_3 = 7.88$	$\tau_3 = -9.67$ $\Phi_2 = 31.17$ $\Phi_3 = 46.75$
BA	$\tau_3 = -3.08$ $\Phi_2 = 3.19$ $\Phi_3 = 4.76$	$\tau_3 = -7.38$ $\Phi_2 = 18.16$ $\Phi_3 = 27.24$
gov	$\tau_3 = -2.99$ $\Phi_2 = 3.03$ $\Phi_3 = 4.54$	$\tau_3 = 9.35$ $\Phi_2 = 29.14$ $\Phi_3 = 43.71$
ZCF	$\tau_3 = -3.03$ $\Phi_2 = 3.33$ $\Phi_3 = 4.97$	$\tau_3 = -8.39$ $\Phi_2 = 23.49$ $\Phi_3 = 35.22$
$ZCIS$	$\tau_3 = -2.72$ $\Phi_2 = 2.70$ $\Phi_3 = 3.76$	$\tau_3 = -8.25$ $\Phi_2 = 22.70$ $\Phi_3 = 34.05$
	$\boldsymbol{\alpha = 1\%}$	$\boldsymbol{\alpha = 5\%}$
Critical values	$\tau_3 = -3.99$ $\Phi_2 = 6.22$ $\Phi_3 = 8.43$	$\tau_3 = -3.43$ $\Phi_2 = 4.75$ $\Phi_3 = 6.49$

Table 5: KPSS test on model variables.

Variable	**Levels Test. stat.**	**Differences Test stat.**
$BEIR$	$\tau = 0.53$	$\tau = 0.03$
CCS	$\tau = 0.42$	$\tau = 0.04$
$v2x$	$\tau = 0.15$	$\tau = 0.02$
BA	$\tau = 0.68$	$\tau = 0.02$
gov	$\tau = 0.44$	$\tau = 0.04$
ZCF	$\tau = 0.33$	$\tau = 0.06$
$ZCIS$	$\tau = 0.88$	$\tau = 0.02$
Critical values	$\boldsymbol{\alpha = 1\%}$ $\tau = 0.216$	$\boldsymbol{\alpha = 5\%}$ $\tau = 0.146$

Table 6: Tests on model residuals.

Test	**Test stat.**	**p-value**
KPSS	0.07	> 0.1
Breusch-Pagan	23.22	0.28
Shapiro-Wilk	0.99	0.70
Ljung-Box	0.28	0.60

B.2 Diagnostic

Figure 6: Regression residuals plot.

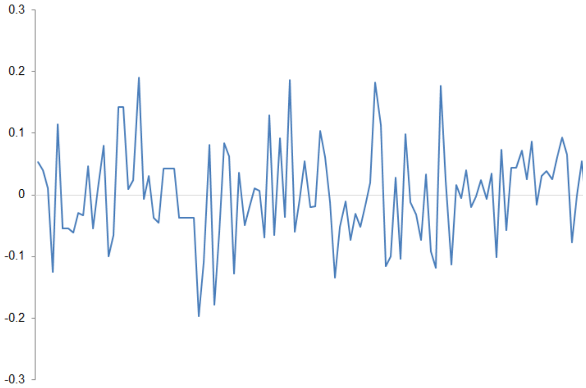
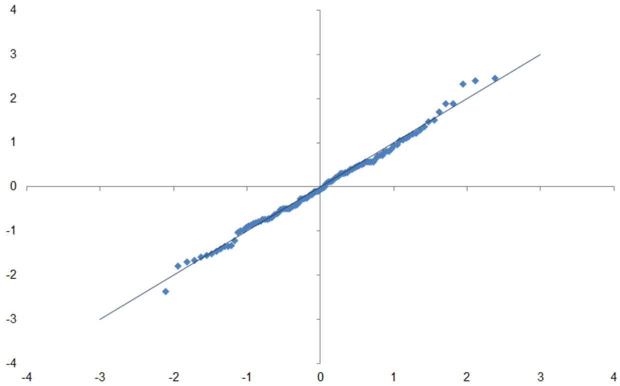


Figure 7: Residuals QQ-plot.



In order to assess the goodness of fit of our model, we report a diagnostic analysis in this section. Starting with a graphical inspection (Figures 6 and 7), the residuals of our regression model appear stationary, homoscedastic, Gaussian and with a negligible serial correlation. In order to verify our qualitative argument from a quantitative perspective, we conducted a test for each of the mentioned characteristics. In particular, we ran (in R) a KPSS test in order to verify stationarity, a Breusch-Pagan test of homoscedasticity, a Shapiro-Wilk test for normality and a Ljung-Box test for the absence of an auto-correlation. The null hypothesis of each of these tests corresponds to what we want to verify; in other words, if residuals are well behaved we have to accept the null hypothesis for every test. The results, which are reported in Table 6, strongly confirm our qualitative considerations.

From the goodness of fit point of view, our model exhibits an adjusted R^2 above the 80% and it fits data well in terms of both level and differences, as can be seen in Figures 8 and 9.

Figure 8: In-sample fit of break-even rate: differences.

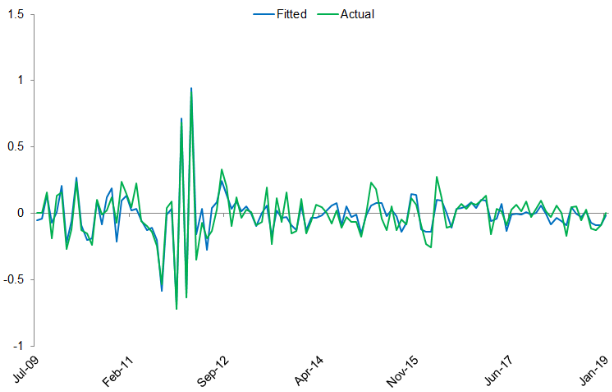
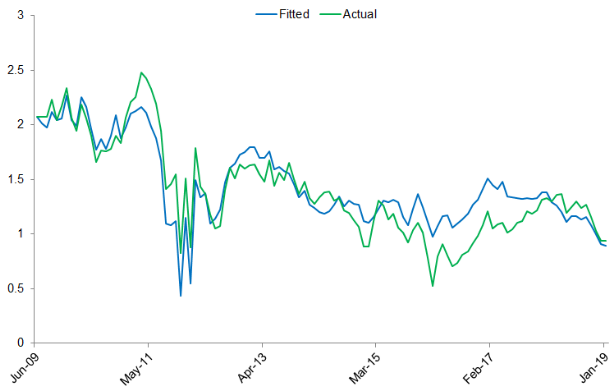


Figure 9: In-sample fit of break-even rate: levels.



Appendix C

Term Structure approach

C.1 Bias Corrected OLS estimates

Figure 10: 5Y5Y inflation expectations.

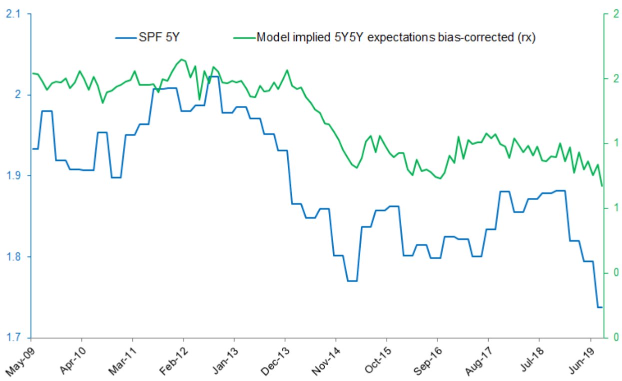
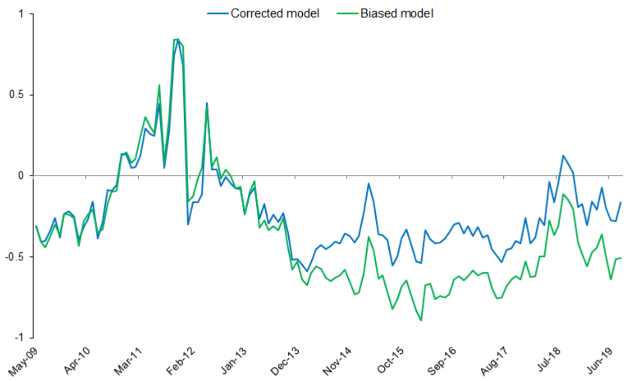


Figure 11: 10-year spot IRP comparison.



In this section we have reported some results obtained by estimating the VAR model of the ASTM and correcting the estimates of the coefficient matrix Φ via bootstrap in order to overcome the persistence bias problem.

As previously stated in Section 4.2, in this work, we have presented and commented on the results of the model without correction principally because it is consistent with survey measures. In fact, when correcting the persistence bias, the estimated (5Y5Y) inflation expectations deviate significantly from the professional forecaster measure. This fact is clearly visible in Figure 10, where the SPF expectations and those obtained using the corrected model are compared. Although the dynamics of the two measures are comparable, with a significant drop starting at the end of 2013 and the descending trend in the last part of the sample period, these two expectations show a huge mismatch on the level of inflation, with a large range of variations in the model-implied inflation. In August 2019, the SPF measure was around 1.75 per cent, while the model-implied 5Y5Y forward inflation is around 1.2 per cent.

In Figure 11 we show the 10-year inflation risk premium estimated with the unadjusted and corrected model. Before 2013, these two estimates are practically indistinguishable, while some differences have started to emerge since 2014. Corrected and biased estimates of the 10-year spot IRP exhibit substantially the same amount of variability and a difference that is only about 13 bp on average; this can be seen as proof of the robustness of our estimates, also using an unadjusted model.

C.2 Diagnostic

Table 7: Yield pricing errors on the whole estimation period. Percentage values.

	Nominal				**Real**			
	2Y	**5Y**	**7Y**	**10Y**	**2Y**	**5Y**	**7Y**	**10Y**
Mean	-0.16	-0.13	0.21	0.02	0.18	-0.09	0.01	0.01
Standard dev.	0.20	0.28	0.25	0.21	0.23	0.13	0.11	0.07
Asymmetry	-2.58	3.32	1.65	-2.68	-0.45	1.28	3.47	1.65
Kurtosis	9.83	16.84	3.60	10.62	-0.42	4.82	18.89	4.55

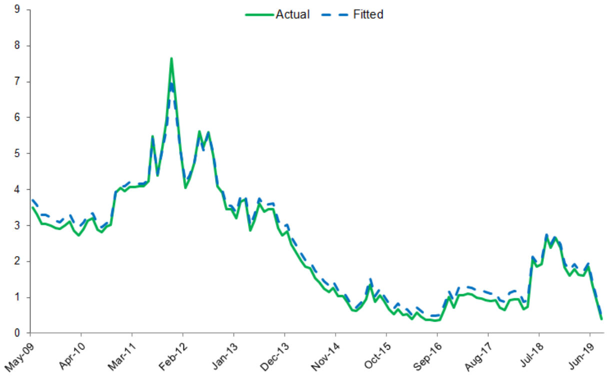
Table 8: Yield pricing errors computed after euro-area sovereign debt crisis (June 2013 - August 2019). Percentage values.

	Nominal				**Real**			
	2Y	**5Y**	**7Y**	**10Y**	**2Y**	**5Y**	**7Y**	**10Y**
Mean	-0.10	-0.16	0.06	0.06	0.26	-0.06	-0.03	-0.01
Standard dev.	0.07	0.05	0.14	0.07	0.25	0.09	0.05	0.04
Asymmtetry	-0.69	0.44	1.18	-0.78	-1.03	0.01	-0.71	-0.15
Kurtosis	-0.77	-0.51	0.12	0.00	0.36	-0.81	0.72	-0.80

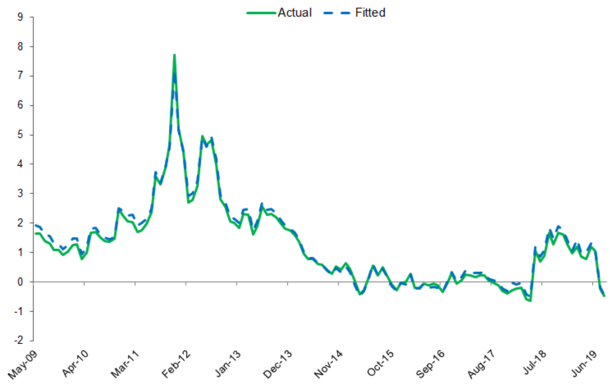
In this section, we show the results of a brief diagnostic analysis on the term structure model. In Tables 7, we have reported pricing errors on yields for the whole sample. The largest errors are obtained on short maturities and especially on nominal yields. Looking at Figure 12, it is clear that we have a lack of fit on yields, principally during the euro-area sovereign debt crisis when yields on Italian bonds spiked. In order to

Figure 12: Fitted versus observed yields. Percentage values, monthly data.

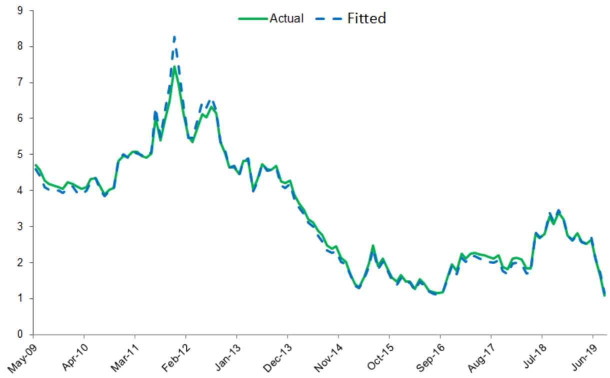
(a) 5-year nominal yield.



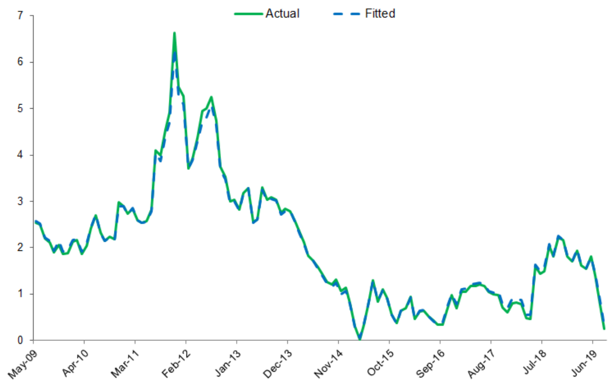
(b) 5-year real yield.



(c) 10-year nominal yield.



(d) 10-year real yield.



assess the goodness of fit of the model net of this effect, in Table 8, we computed yield pricing errors only considering the post-crisis period (starting from June 2013). Although pricing errors in this shortened sample are lower than the ones in the whole sample, there remains some lack of fit on short maturities bonds. We stress here that our focus is on medium/long-term bonds where the errors are in line with other studies.