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Investor strategies in the green bond market: the influence of liquidity risks, economic factors and clientele effects

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Abstract

The green bond market has dramatically expanded especially in Europe but severe liquidity issues may undermine its rapid development. If few studies have assessed the implied liquidity risks for investors in terms of liquidity premium, none of them have specifically analysed its behavior across bond maturities. To fill this gap, this paper studies the term structure of the liquidity premium of the green bond market.

We find that the sizes of short-term and long-term premia are close to those estimated on the German government bond market. We show that those premia are affected by economic factors and by spillover effects between them, which contribute to the U-Shape of the liquidity premium. Finally, we detect a liquidity clientele effect on the ask side impacting the liquidity premium, which implies a maturity segmentation i.e., high-risk (*resp.* low-risk) investors buy short-term (*resp.* long-term) green bonds and hold them until maturity.

Taken together, our results deliver valuable insights on investors' strategies in the green bond market. Quite importantly, green bond investors prefer to opt for buy and hold strategies because they are compensated for higher liquidity risks along the entire maturity spectrum.

Keywords: Green Bond ; Liquidity Premium ; Term structure ; Spillover effects ; Clientele effect

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I. Introduction

A major and recent breakthrough in sustainable finance is the issuance of green bonds. These bonds are intended to play an extensive role in financing climate friendly projects with long-term investment horizons and large capital costs (e.g., renewable energy) or related to urgent environmental transitions (e.g., water management) (Tolliver et al., 2020 ; CBI, 2021).

In 2020, the global green bond primary market amounts to USD 290 bn of issuances where the European market concentrates 48% of this total (CBI, 2021). Interestingly, this market emerged in the midst of the 2008 financial crisis, a liquidity driven crisis culminating in a credit crunch and global recession. General concerns about bond liquidity have persisted ever since due to tighter regulation and important costs of technology development (EC, 2017). Against this backdrop, green bonds have been able to attract greater attention of investors, notably an investor clientele searching for environmental and long-term benefits (Flammer, 2021). However, the shortage of green bond issuances implies a low liquidity reflected into wider bid-ask spreads and higher liquidity risks (Bachelet et al., 2019; Fender et al., 2019). Another source of liquidity risk is the lack of reporting transparency on the use of green bond proceeds resulting in important adverse selection costs (Febi et al., 2018).

This liquidity risk is a relevant topic for investors and issuers for at least two reasons. First, it raises implied yields through a liquidity premium which compensates investors for carrying (il)liquidity risks (Bao and al., 2011). If it is sufficiently high to reward for the liquidity risk of green bonds, investors are more likely to buy and hold them rather than trading them to dynamically manage (rebalancing) bond portfolios. Second, it is a relevant metric for issuers that usually have multiple bonds outstanding. Green bonds are close substitutes with respect to credit risk and mostly differ in their maturity (Zerbib, 2019). In this respect, bond issuers can target a specific investor base or liquidity clientele i.e., high-risk (*resp.* low-risk) investors when they decide to issue short-term (*resp.* long-term) green bonds. This is not without incidence on their financing costs, and their capital structure (N’Guyen et al., 2021). These two reasons point to using the term structure of liquidity premia as a natural way to study the impact of the liquidity risk on green bond yields. Finally, from a risk management perspective, if some economic and/or risk factors determine this term structure, relevant hedging instruments differ across maturity segments accordingly (Kempf et al., 2012).

Although the existence of the liquidity premium for green bonds is beyond doubt, we remain unclear about its size, its term structure and its main determinants. To fill this lacuna, this paper aims at (1) estimating the size of the liquidity premium across bond maturities and (2) studying how specific investor strategies along economic factors influence its term structure.

The green bond market offers a unique testing ground for liquidity studies: Highly rated bonds that only differ with respect to their green credentials and to liquidity are traded along the entire maturity spectrum. In this paper, we take advantage of this situation to examine the size and the term structure of liquidity premia from a panel of about 200 green bonds issued in Europe with tenors between 2 to 20 years over a five-year period. We proceed in two steps.

First, we adapt the liquidity regression model developed by Dick Nielsen et al. (2012) to quantify the liquidity premium. We use the bid-ask spread to identify the contribution of liquidity to bond yields, across rating categories and maturity. Within a rating category, we order the bonds according to their bid-ask spreads, compute the difference between the 5% and the 50% quantiles and multiply by the regression coefficient for that rating category. We benchmark these measures with those estimated from the difference between the median value of Green Bonds issued by Supranationals, Sovereigns and Agencies (SSAGB hereafter) and those issued by Corporates (CGB hereafter). The large difference in liquidity between these two market segments provides a convenient setting to quantify the liquidity premium.

Second, we estimate the term structure of the liquidity premium using the Nelson and Siegel (NS) (1987) model, which allows us to divide it in three parts: short-, medium- and long-term. Using the NS framework in our analysis presents two other immediate advantages. On the one hand, the estimated liquidity premium can be viewed as a robust measure of liquidity as it captures the forward-looking component of liquidity (i.e., investors' expectations on the future market liquidity). On the other hand, using the NS model to estimate the liquidity premium is not new (e.g., Kempf et al., 2012; Schuster and Uhrig-Homburg, 2015) so we can compare our estimates to those obtained by these empirical studies.

Given this methodological framework, our paper is the first, to the best of our knowledge, to address important questions related to the liquidity premium of the green bond market: What is its size across different bond maturities and the shape of its term structure? Is its term structure influenced by economic factors? Is it also driven by specific maturity bond effects such as spillover, liquidity clientele and aging effects identified in mature bond markets?

To answer these questions, we build on the insights of the literature on bond markets.

At an empirical level, Longstaff (2004) measured a liquidity premium ranging from 10 to 15 bps on the U.S. Treasury bond market. Schuster and Uhrig-Homburg (2015) estimated the liquidity premium of the German government bond at 15 bps by using the KfW-Bund yield spread. This figure is higher than this found by Kempf et al. (2012), who estimated a long-term (*resp.* short-term) premium of 12 bps (*resp.* 10 bps) for the German government bond market by comparing the yields of the liquid Bunds and those of the less liquid Pfandbrief bonds.

Regarding the term structure of the liquidity premium, there is little consensus among empirical studies, which conclude that it is either decreasing (e.g., Chen et al., 2020), increasing (e.g., Dick-Nielsen et al., 2012), or even U-shaped (e.g., Schuster and Uhrig-Homburg, 2015). A large body of literature argues that short and long dated bonds are not perfect substitutes since their (liquidity) investor clienteles differ. Gehde-Trapp et al. (2018) verify two important results related to investor clienteles: (i) investor portfolio decisions involve a transmission of liquidity shocks from the short end to the long end of the term structure, generating spillover effects; (ii) a U-shaped liquidity premium, where short and long maturities carry the largest compensation for illiquidity risks. This latest result implies that issuers can choose an investor clientele: low-risk (*resp.* high-risk) investors when issuing short-term (*resp.* long-term) bonds.

Our paper contributes to the above-mentioned literature in three important respects.

Our first contribution is related to the term structure of the liquidity premium of the green bond market. We report that the long-term liquidity premium of the green bond market is larger (16 bps) than the short-term one (13 bps) on average. These values are close to those estimated on the German government bond market by Schuster and Uhrig-Homburg (2015). However, the difference between the short-term and long-term liquidity premia is noticeably lower than this found by Dick-Nielsen et al. (2012) on the corporate bond market. This slight difference notably results from spillover effects running from the short end to the long end of the liquidity premium and vice-versa as in Gehde et Trappe et al. (2018), which are found to be more pronounced from 2018. Combined with an on-the-run effect, these spillover effects contribute to the U-shape of the liquidity premium of the green bond market.

Our second contribution lies in the detection of a liquidity clientele effect on the ask side, which increases the ask liquidity premium (Goyenko et al., 2011). This effect favours an insufficient selling pressure, i.e., sales at discounted prices by liquidity-shocked investors because they are prone to buy green bonds and hold until maturity (Dick-Nielsen et al., 2012). Also, this clientele effect implies segmentation among those investors (Beber et al., 2021) given that issuers may target an investor base i.e., high-risk (*resp.* low-risk) investors when issuing short-term (*resp.* long-term) green bonds. This choice of an investor base impact their financing costs and their capital structure (N'Guyen et al., 2021). Second, we show that this clientele effect is only related to the liquidity channel but is not due to the credit risk channel. Overall, these two important findings support implications of the clientele effect, which are both the differential impact on short-term and long-term bonds and the absence of effect of funding constraints on bond pricing consistent with Chen et al. (2020).

Our third contribution is related to the impact of economic factors on the liquidity premium of the green bond market. We show that the economic outlook and the economic policy uncertainty determine both its short-term and long-term liquidity premia moderately as is the case with the German government bond market (Schuster and Uhrig-Homburg, 2015). In addition, the performance of the U.S Municipal tax-exempted green bond market positively affect the long-term liquidity premium. Said differently, applying tax exemptions to the coupons of green bonds decrease the long-term liquidity premium so the compensation of long-term investors for higher (il)liquidity risks when they buy and hold green bonds until maturity.

Taken together, our results underline the importance of the liquidity risk for investors when they decide to trade green bonds for pro-environmental choices, tax incentives (Zerbib, 2019), diversification (Sangiorgi and Schopohl, 2021) or carbon hedging (Jin et al., 2020; Rannou et al., 2021).

The rest of the paper is organised as follows. Section 2 presents the relevant literature and hypotheses related to the size and the term structure of the liquidity premium of the green bond market. Section 3 describes the data and the methodology used. Section 4 outlines empirical results related to the liquidity premium and Section 5 discusses its drivers. Section 6 concludes.

II. Literature Review and Hypothesis Development

2.1. Key features of the Green Bond market

Green bonds are debt instruments issued by capital raising entities to fund their environmentally friendly projects. These projects may have long-term investment horizons, large capital costs, secured cash flows like renewable energy infrastructures (Tolliver et al., 2020) or may relate to urgent environmental transitions (e.g., sustainable water management). They have all the features of ordinary bonds with the additional green characteristic designating that their “use of proceeds” must be geared towards positive environmental outcomes.

Since the first corporate green bond was issued in 2013, Europe has led the global green bond market, representing 48% of total issuances in 2020 (USD 156bn), where supranational institutions (e.g., World Bank, EIB), financial and non-financial corporates have each contributed to 25% of the total (CBI, 2021). In Europe, the Luxembourg Green Exchange (LGX) has overseen the rapid development of the green bond market. From Table 1, we can see that LGX is the most active platform in terms of issuance, followed by LSEG and Euronext.

[INSERT TABLE 1 HERE]

2.2. *The importance of liquidity risk on the green bond market*

Demand for green bonds has initially been driven by institutional investors with ESG mandates (e.g., Allianz Global Investors, Mirova) and dedicated green bond funds (e.g., State Street Calvert, BlackRock). More recently, a broader range of investors including asset managers, insurance companies, banks, corporate treasuries, retail investors, and the European Central Bank invested the European green bond market. With the advent of green bond indexes and green bond exchange-traded funds (ETFs), green bond investing becomes part of passive portfolio strategies or used as carbon hedging strategies (Jin et al., 2020; Rannou et al., 2021).

Analysing a survey conducted by CBI in 2019 allows Sangiorgi and Schopohl (2021) to note that asset managers consider liquidity as critical when investing in green bonds since they rank the criteria of minimum size and liquidity of the bond (3.938/5) given a Likert scale from one (lowest score) to five (highest score). Bachelet et al. (2019) report that the difference in bid-ask spreads between green bonds issued by institutional (*resp.* private) issuers and their ordinary twins is 10 bps (*resp.* 18 bps). For Bachelet et al. (2019), such a difference indicates higher (il)liquidity risk for green bonds, which are priced in yields (Zerbib, 2019).

2.3. *Hypothesis development*

Having studied a limited sample of green bonds issued between 2013 and 2017, Febi et al. (2018) found that both their bid-ask spreads and LOT measures were positively related to yield spreads confirming that the (il)liquidity risk is priced in green bond yields. They estimated an impact of LOT liquidity premium on green bond yield spreads twice higher than that of the US investment grade corporate bonds (Chen et al., 2007). However, there is, to our knowledge, no attempt to estimate the size of the liquidity premium of the green bond market.

In the following paragraphs, we build on the insights of the existing literature on bond markets to derive three testable hypotheses which are used to outline the relevant aspects of the liquidity premium of the green bond market across maturities and its main determinants.

Schuster and Uhrig-Homburg (2015) estimated the liquidity premium of the German government bond market at 15 bps when Kempf et al. (2012) measured long-term (*resp.* short-term) premia at 12 bps (*resp.* 10 bps) for the same market on average. Besides, there is less consensus about the shape of the term structure of the liquidity premium: decreasing (e.g., Chen et al., 2020), increasing (e.g., Dick-Nielsen et al., 2012), or U-shaped (e.g., Schuster and Uhrig-Homburg, 2015). For Dick-Nielsen et al. (2012), the liquidity premium is twice as high for long-dated bonds compared to short-dated bonds while for Chen et al. (2020), long-dated U.S. corporate bonds have lower liquidity premia. By contrast, Schuster and Uhrig-Homburg (2015)

documented a U-Shape pattern for the liquidity premium on the German government bond market. These considerations lead to the following testable hypothesis **H₁**.

H₁. The liquidity premium of the green bond market is similar to this observed on the German government bond market in terms of size and term structure.

The literature on bonds has also provided evidence of the influence of economic factors on the liquidity premium. On the German government bond market, Kempf et al. (2012) show that investors' expectations about the long-term economic outlook determine the long end of the liquidity premium. On the same market, Schuster and Uhrig-Homburg (2015) find that the sensitivity of the liquidity premium to economic fundamentals rises in stressful economic periods. On the green bond market, Broadstock and Cheng (2019) show that macroeconomic factors such as economic policy uncertainty influence its development. Previously, this economic policy uncertainty was found to affect the long-term liquidity premium (Kempf et al., 2012). Given these different findings, we consider the following testable hypothesis **H₂**:

H₂. The term structure of the liquidity premium of the green bond market is influenced by economic factors notably economic outlook and economic policy uncertainty.

Gehde-Trappe et al. (2018) found that investor-portfolio decisions induce a transmission of liquidity shocks from short maturities to long ones generating spillover effects. Instead, Kempf et al. (2012) do not detect spillover effects on the German government bond market. In addition to those effects, the age of the bond, more precisely, the on-the-run vs. off-the-run effects (see Mahanti et al. (2008) for an in-depth explanation) may affect the liquidity premium. The rationale is that when a bond is recently issued, it is on-the-run, enjoying higher liquidity and paying lower liquidity premium compared to when it has been outstanding for a while i.e., off-the-run. De Jong et Driessen (2015) note that studies often confirmed a lower (*resp.* greater) liquidity premium for the on-the-run (*resp.* off-the-run) treasury or government bonds.

Given all of these findings, we formulate **H₃** in order to test the influence of spillover effects and aging effects (i.e., the on-the-run vs off-the run effect) on the liquidity premium.

H₃. Both spillover and aging effects affect the shape of the term structure of the liquidity premium of the green bond market.

Gehde-Trapp et al. (2018) verify a U-shape of the liquidity premium on the corporate bond market, where short and long maturities carry the largest compensation for illiquidity risks. They explained that this pattern results from a liquidity clientele effect, i.e., the tendency for long-term (*resp.* short-term) investors to hold bonds with high (*resp.* low) illiquidity risk.² Conversely, Amihud and Mendelson (1986) assuming constant bid-ask spread demonstrate that the liquidity clientele effect implies a decreasing term structure of the liquidity premium. Given these considerations, we consider **H4** to test the potential impact of a liquidity clientele effect on the term structure of the liquidity premium.

H4. A liquidity clientele effect impacts the term structure of the liquidity premium of the green bond market and investor strategies.

III. Data and Methodology

3.1. Data selection

We constitute a representative sample of green bonds issued on LGX, the most active exchange for green bond listings in Europe, which attracts a diverse and international clientele of fund managers, retail and institutional investors. Our sample is composed of fixed rate bullet green bonds, which have been issued from 2011, with more than 100,000 euros (or equivalent) initial amount, and with a maturity between one year and twenty years. We exclude therefore floating rate bonds, and index linked bonds, as well as 144A bonds for liquidity purposes.³

After applying additional filters, we obtain a final sample of 184 green bonds that have been issued by 51 entities since 2015 because most of corporate green bonds were issued after 2014.⁴ Our final sample is composed of 91 CGBs and 93 SSAGBs that we study over an extensive 5-year period: January 2015 - December 2019 through their daily closing prices and bid/ask quotes. All closing (clean) prices are provided by the LGX exchange. Bid and ask quotations are collected from Bloomberg and are used to estimate bid-ask spreads.

3.2. Descriptive statistics

Panel A of Table 2 shows summary statistics for the whole sample and for the two subsamples: SSAGBs and CGBs. SSAGBs have slightly longer average maturities that can

² Even if short-term (*resp.* long-term) bonds are not more (*resp.* less) liquid given the bid-ask spread, they are more (*resp.* less) liquid due to the lower (*resp.* higher) disutility from awaiting their closer (*resp.* more distant) maturity.

³ Rule 144A bonds, may not have any liquid secondary market or ready purchaser since they are limited to trading among qualified institutional investors. We remove Floating Rate Notes (FRNs) and index-linked bonds from our initial sample as it was done by the empirical literature (see e.g., Bachelet et al., 2019). In fact, their coupons are variable making impossible an accurate calculation of their relative bond values and their corresponding yields.

⁴ We exclude bonds that have less than one year of data (prices or quotes), implying that green bonds must be issued before December 31, 2019 or those for which the bid and ask quotations are not available or equal.

explain that they pay a higher coupon. They have also a credit rating close to AAA while CGBs are at mid distance between an AAA and AA rating (based on S&P's rating). We also notice that green bonds issued by corporates involve higher yields than those of SSA while they appear less liquid as in Bachelet et al. (2019) consistent with the idea that liquidity is priced in bond yields (Zerbib, 2019). Indeed, the average issuance size of CGBs is 0.7 times lower than this of SSAGBs suggesting that CGBs are less liquid. To confirm the significant difference in terms of liquidity between the CGB and SSAGB segments, we use three complementary proxies: the closing bid-ask spread, consistent with Zerbib (2019), the zero trading days (ZTD) like Febi et al. (2018) and Bachelet et al. (2019), and the Roll measure as in Dick-Nielsen et al. (2012).

The bid-ask spread is arguably the most utilised measure of liquidity costs and has been widely used as a major proxy of the degree of green bond liquidity (e.g., Fender et al., 2019; Lebellet et al., 2021). As for ZTD, it is used as a liquidity proxy in a host of empirical studies (Chen et al., 2007). Similar to Febi et al. (2018), we measure ZTD as the proportion of days with zero returns within a month assuming that the daily return is zero when the market is inactive and illiquid (i.e., no trading activity leads to the absence of daily price changes). The Roll measure of price reversals is based on the covariance between subsequent returns. In contrast with Febi et al. (2018), we do not compute the LOT measure (Chen et al., 2007) since it is based on an unrealistic assumption of a normal distribution for green bond returns.

Panel B of Table 2 presents distributions statistics for the three liquidity proxies for the whole sample of green bonds. The average bid-ask spreads of SSAGBs are 20% tighter than those of CGBs. This difference is significantly greater if we consider the ZTD and the Roll measure for CGBs that are twice as high on average. The median number of ZTD is 5%, in line with the estimates of Bachelet et al. (2019) confirming that the green bond market is relatively an illiquid market compared to other corporate bond markets. If the median percentage BAS is a rather low figure (0.322%), BAS is comprised in a wide range. The median Roll measure is slightly higher than this found on the corporate bond market by Dick-Nielsen et al. (2012).

Overall, two key results emerge from Table 2 namely (i) green bonds issued by corporates (CGBs) are clearly less liquid than those issued by Supranationals, Sovereigns and Agencies (SSAGBs) owing to higher bid-ask spreads and smaller issuance size (see Mahanti et al., 2008) and (ii) CGBs and SSAGBS have similar credit risk profiles. These two findings suggest that the liquidity differential between CGBs and SSAGBs may be used to adequately capture the value of the (intramarket) liquidity premium that prevails in the green bond market.

[INSERT TABLE 2 HERE]

3.3. Liquidity regressions

To get a first idea on the influence of liquidity, we regress bond yields on liquidity proxies one at a time while controlling for credit risk similar to Dick Nielsen et al. (2012) in Eq. (1) :

$$Yield_{i,t} = \alpha + \gamma_1 \cdot Liquidity_{i,t} + \beta_1 \cdot Bond\ age_{i,t} + \beta_2 \cdot Amount\ issued_{i,t} + \beta_3 \cdot Coupon_{i,t} + \beta_4 \cdot TTM_{i,t} + \beta_5 \cdot Ratingscale_{i,t} + \beta_6 \cdot Term\ slope_{i,t} + \beta_7 \cdot Credit\ spread + \beta_8 \cdot Certified + \beta_9 \cdot Currency_{i,t} + \varepsilon_{i,t} \quad (1)$$

Where: i is the bond issue, t is the month, $Yield_{it}$ is the bond yield at the end of t , $Liquidity_{it}$ is the liquidity proxy i.e., the percentage BAS, ZTD or the Roll measure, $BondAge$ is the age of the bond, $Amount\ issued$ is its amount issued in euros, $Coupon$ is its coupon rate, TTM is the time to (bond) maturity, $Ratingscale$ is its rating category based on S&P rating equivalence, $Term\ slope$ the difference between the 10-year and 2-year German government bonds' rate, $Credit\ spread$ is the credit risk on the corporate bond market, $Certified$ if the bond is CBI labelled and/or has been externally certified, $Currency$ is the currency of the bond.

Since we have panel data set of yields with issuers potentially having more than one bond outstanding at any point in time, we calculate two-dimensional cluster robust standard errors consistent with Dick-Nielsen et al. (2012). This corrects for time series effects, firm fixed effects and heteroscedasticity in the residuals.

We use bid-ask spreads to proxy *Liquidity* since it the most central variable used to explain liquidity premiums (Schuster and Uhrig-Homburg, 2015) and alternatively ZTD or Roll measure as in Dick-Nielsen et al. (2012). As control variables, we use intrinsic bond characteristics through which green bonds differ. The variables considered are the age of the bond (*Bond age*), *Currency*, the time-to maturity (*TTM*), the coupon rate (*Coupon*) and the issuance amount of the green bond (*Amount issued*) consistent with Dick-Nielsen et al. (2012). *Ratingscale* assesses the issuer' credit risk on liquidity based on S&P rating equivalence (Febi et al., 2018; Bachelet et al., 2019). As in Kempf et al. (2012), we include a *Credit spread* variable in the regression, i.e., the spread between the Bloomberg EUR Eurozone index of AA+/AA bond yields and the Bloomberg EUR Eurozone index of BBB bond yields to capture the impact of credit risks on the European corporate bond market. Similarly, we consider *Term slope*, i.e., the difference between the 10-year and 2-year German government bonds' rate to capture the level of credit risk on the European government bond market like Febi et al. (2018).

Finally, we define a dummy *Certified* variable as in Bachelet et al. (2019). It takes the value of 1 (0 otherwise) when the bond meets the CBI requirements and/or has been externally certified by an external reviewer. The rationale is that green bonds with CBI label or externally certified carry less adverse selection, so the liquidity premium is lower (Febi et al., 2018).

Panels A of Table 3 shows that BAS coefficients on bond yields are significant at 1% level, so the evidence that liquidity costs are priced is conclusive, consistent with the finding in Chen et al. (2007). This indicates that BAS tend to increase bond yields whatever the rating is. Running regressions for different rating categories shows how robust our conclusions are regarding the effect of liquidity. We do this for three rating categories: AAA, AA, A and –. We also see that the Roll measure has positive coefficients across all ratings and most of them are statistically significant at 1% level. Turning to zero-trading days (ZTD), we find that there is consistent positive relationship between ZTD and bond yields. The link between ZTD and bond yields is consistent with the results of Bachelet et al. (2019). Panels B and C of Table 3 present the coefficients of liquidity proxies when ask yields and bid yields as dependent variables are used. They are found to be similar to those observed in Panel A both in magnitude and in sign.

[INSERT TABLE 3 HERE]

Besides, we report in Appendix A the values of all coefficients presented in Eq. (1) using the full sample of green bonds and our three liquidity proxies (see Table A1). We can see that all variables capturing the influence of credit risks (*Ratingscale*, *Term slope*, *Credit spread*) are not significant at both 1% and 5% levels in all cases. This result clearly indicates that the credit risk is not significantly priced in green bond yields in contrast with the liquidity risk. Instead, we find that coefficients of the variable *Certified* are significant at 5% level, which suggests that asymmetry information costs affect green bond yields as shown by Febi et al. (2018). Finally, we verify that coefficients of *Bondage* are always significant at 1% level confirming the relevance to study the behavior of yields and liquidity costs across different maturities.

IV. Liquidity premium

4.1. Size of the liquidity premium

In order to calculate the impact of green bond (il)liquidity on their yields, we run the following pooled regression based on two rating categories denoted R: AAA and AA and -.

$$Yield_{i,t}^R = \alpha^R + \gamma_1^R \cdot Liquidity_{i,t} + \beta_1^R \cdot Bond\ age_{i,t} + \beta_2^R \cdot Amount\ issued_{i,t} + \beta_2^R \cdot Coupon_{i,t} + \beta_3^R \cdot TTM_{i,t} + \beta_4^R \cdot Term\ slope_{i,t} + \beta_5^R \cdot Credit\ spread + \beta_7^R \cdot Certified + \beta_8^R \cdot Currency_{i,t} + \varepsilon_{i,t} \quad (2)$$

Where: i is the bond issue, t is the month, $Yield_{i,t}^R$ is the bond yield calculated at the end of t , $Liquidity_{i,t}^R$ is the liquidity proxy i.e., the percentage BAS, ZTD or the Roll measure.

All other control variables are also in Eq. (1) and are defined in the above paragraph.

Table 4 displays the estimated size of the *Liquidity* component and the confidence bands calculated from a wild cluster bootstrap of the regression residuals such as in Dick-Nielsen et al. (2012). Panel A presents the estimates obtained from a method developed by Dick and Nielsen et al. (2012). Within the two rating categories (AAA, AA and -), and three maturity ranges (2–5 years, 5–10 years, 10–20 years), we sort all observations according to their bid ask spread (liquidity) measure. The liquidity component is defined as the difference between the 50% quantile and the 5% quantile of the bid-ask spread distribution. It therefore measures the difference in bond yields between a bond with average liquidity and a very liquid bond.⁵ For AAA rated bonds, the component is on average 3 bps higher than the non-default component of the German government bond market estimated by Schuster and Uhrig-Homburg (2015). We see that the liquidity component barely increases as the rating decreases. For AA and – rated bonds, the liquidity component is estimated to be 1 bp higher than this of AAA rated bonds. This result shows that the credit risk profile of green bonds does not really impact the three liquidity premia (Febi et al., 2018). It also indicates an absence of a flight-to-quality phenomenon leading investors into buying top rated bonds regardless of their liquidity.

Panel B shows the size of the liquidity component based on the liquidity differential between the CGB and SSAGB segments, which is likely to capture the “intramarket” liquidity premium. Consequently, the liquidity premium is defined as the difference between the median value of the bid-ask spread for CGBs and this for SSAGBs respectively. Its size estimated in AAA bonds (*resp.* AA and – bonds) is 13.7 bps (*resp.* 14.4 bps), which is comparable in magnitude to the liquidity premiums of the German government bond market (Schuster and Uhrig-Homburg, 2015) and of the US treasury market (Longstaff, 2004).

Remarkably, the liquidity premium appears to be a positive function of the maturity for the two rating categories. The premium of long-dated bonds (10-20 years) is between 25% (Panel A) and 30% (Panel B) higher than this of short-dated bonds (2-5 years). This difference is similar to this found by Longstaff (2004) on the US treasury bond market and Kempf et al. (2012) on the German government bond market. It can be explained by the fact that the liquidity premium captures different aspects of illiquidity including the liquidity risk in addition to informational risks and search for counterparty costs (see Dick-Nielsen et al., 2012).

[INSERT TABLE 4 HERE]

In Appendix B, we carry out robustness tests related to Table 3 results. Specifically, we estimate the liquidity premiums from the coefficients of Roll measure obtained in Eq. (2).

⁵ For comparison purposes, we consider two rating categories (AAA ; AA and -) in Panels A and B since our sample does not contain SSAGBs with a rating lower than AA.

Panels A and B of Table B1 show similar results in terms of size across maturities using the differential between SSAGBs and CGBs whereas the method based on Dick-Nielsen et al. (2012) provides significantly lower estimations compared to Table 3 and wider confidence intervals. This supports our decision to choose the liquidity differential between SSAGBs and CGBs as a more consistent basis to study the dynamics of the liquidity premia further.

Overall, the results of Table 4 and Table B1 validate H_1 provided that the sizes of the liquidity premia of the green bond market across different maturities are close to those observed on the mature German government bond market.

We then compute the fraction of the liquidity component to the bond yields.

Panels A and B of Table 5 are based on the liquidity components estimated in Table 3 according to the method of Dick Nielsen et al. (2012) and to the liquidity differential between CGBs and SSAGBs respectively. A consistent finding from Table 3 is that the liquidity portions of yields in AAA rated bonds are close to those obtained for AA rated bonds and lower grades confirming the absence of a flight-to-quality effect towards prime quality green bonds.

Panels B presents lower liquidity portions of the yield for short maturities (between 10% and 10.5% for the maturity range 2-5 years) compared to those for long maturities (between 12 and 12.5% for the maturity range 10-20 years). This result is also observed in Panel A, where higher fractions for all maturities and all ratings (16% or less) are reported. Interestingly, these fractions are at lowest levels for the maturity bucket: 5-10 years for Panels A and B. This slight decline at the intermediate maturity bucket may be either related to an on-the-run effect (Dick-Nielsen et al., 2012) or to a spillover effect (Gehde-Trappe et al., 2018). Further investigation of this result is deemed necessary and will be carried out in the following paragraphs.

[INSERT TABLE 5 HERE]

4.2. The term structure of the liquidity premium

Table 3 only offers a static analysis on three liquidity premia (short: 2-5 years; intermediate: 5-10 years, and long: 10-20 years). Hence, we propose to study the dynamics of those premia over the period 2015-2019. For that purpose, we estimate a Nelson-Siegel (NS) (1987) model to determine a fitted curve consistent with the approach of Kempf et al. (2012) and Schuster and Uhrig-Homburg (2015). This approach is preferred over the Svensson (1994) model to avoid overfitting and facilitate the interpretation.⁶

⁶ De Pooter (2007) lists three main reasons for which the NS (1987) model is a popular term structure estimation method. First, it provides a parsimonious approximation of the yield curve using only three parameters (contrary to spline methods) that give the model enough flexibility to capture either monotonic, humped or S-type shapes observed in yield curve data. Second, it produces yield curves which have the desirable property of starting off

Within a NS framework, we estimate a zero-coupon yield of bond class $i \in \{CGB, SSAGB\}$ at time t for time to maturity T given by the below equation:

$$\text{Yield}_t^i(T) = \beta_{0,t}^i + \beta_{1,t}^i \left[\frac{1 - e^{-T/\tau_t^i}}{T/\tau_t^i} \right] + \beta_{2,t}^i \left[\frac{1 - e^{-T/\tau_t^i}}{T/\tau_t^i} - e^{-T/\tau_t^i} \right] \quad (3)$$

In order to estimate the four NS parameters ($\beta_{0,t}$, $\beta_{1,t}$, $\beta_{2,t}$, τ_t) for the SSAGB and CGBs segments, we select end-of-month yields with maturities of 2, 3, 4, 5, 6, 7, 8, 9, 10, 12, 15 and 20 years for each bond class. Further, we run the estimation by minimizing the sum of squared yield differences over all selected maturities and for each bond class.

To make the NS estimated parameters comparable, we assume τ_t to be constant over time and identical for the two bond classes i.e., $\tau_t^{\text{CGB}} = \tau_t^{\text{SSAGB}} = \tau_t$ as it is a common practice (see Diebold and Li, 2006). This procedure provides parameter estimates for the CGB class ($\beta_{0,t}^{\text{CGB}}; \beta_{1,t}^{\text{CGB}}; \beta_{2,t}^{\text{CGB}}$) and for the SSAGB class ($\beta_{0,t}^{\text{SSAGB}}; \beta_{1,t}^{\text{SSAGB}}; \beta_{2,t}^{\text{SSAGB}}$) with a mean estimate of $\tau = 2.98$ generating a maximum factor loading of $\beta_{2,t}$ at about 6 years to maturity. Thus, we obtain a mean absolute fitting error of 2.75 bps (*resp.* 2.85 bps) for the SSAGB (*resp.* the CGB) class.

We now interpret the NS parameters: $\beta_{0,t}$, $\beta_{1,t}$, $\beta_{2,t}$ in terms of factors. $\beta_{0,t}$ is the long-term factor while $\beta_{2,t}$ is the shape factor that determines medium-term yields. The short-term factor is represented by $\beta_{0,t} + \beta_{1,t}$ that commands short-term yields (Diebold and Li, 2016). As a result, the differences: $\text{Long}_t \equiv \beta_{0,t}^{\text{CGB}} - \beta_{0,t}^{\text{SSAGB}}$, $\text{Short}_t \equiv \beta_{0,t}^{\text{CGB}} - \beta_{0,t}^{\text{SSAGB}} + \beta_{1,t}^{\text{CGB}} - \beta_{1,t}^{\text{SSAGB}}$, $\text{Shape}_t \equiv \beta_{1,t}^{\text{CGB}} - \beta_{1,t}^{\text{SSAGB}}$ determine the short, medium and long parts of the liquidity premium respectively.

Fig. 1 shows that the global evolutions of the three factors related to the SSAGB and CGB segments are highly correlated. Besides, we observe that the positive long-term and short-term factors of the CGB segment are significantly above those of the SSAGB segment, inducing large short-term and long-term liquidity premia. Also, the negative shape factors of CGBs are always lower than the shape factors of SSAGBs (i.e., close to zero) leading to lower the liquidity premium for medium maturities compared to short-term and long-term liquidity premia. Taken together, these two differences involve a U-shape term structure of the liquidity premium as detected by Kempf et al. (2012) on the German government bond market consistent with **H1**.

[INSERT FIGURE 1 HERE]

Fig. 2 shows that the long-term liquidity premium is always larger than the short-term one suggesting that long-term investors are rewarded for higher illiquidity risks (Bao et al., 2011). On average, the long-term liquidity premium is 1.2 times larger (15.4 bps) than the short-

from an instantaneous short rate value of $\beta_{1,t} + \beta_{2,t}$ and levelling off at a finite infinite-maturity constant value of $\beta_{1,t}$. Third, the three NS components have a clear interpretation as short, medium and long-term components.

term one (12.2 bps). These estimates are close to those found on the German government bond market by Schuster and Uhrig-Homburg (2015) but slightly higher than those found on the same bond market by Kempf et al. (2012) in accordance with **H1**. Both short-term and long-term liquidity premia increase until mid-2019 at a comparable pace and slightly decrease after. Consequently, the correlation between these premia is high (0.73) and significant as is the case with the German government bond market (Kempf et al., 2012). The slight dip in the short-term liquidity premium from mid-2019 may be due to the increasing trading activity of ETF indexers. Indeed, green bond ETFs are likely to be more liquid than their underlying green bonds, attracting shorter-term investors and stimulating short-term demand for liquidity, which may reduce the short-term liquidity premium (Broman and Schum, 2017; Dannhauser, 2017).

[INSERT FIGURE 2 HERE]

Finally, in order to provide further support that the liquidity premia are driven by liquidity differences between CGBs and SSAGBs, we plot the evolution of liquidity premia together with the differences in bid-ask spread across two maturities: 3-Year (short-term) and 10-Year (long-term) in Fig. 3. We note that the bid-ask spread evolutions mimic those of the short and long-term liquidity premia. Also, their large unconditional correlations of 0.69 (*resp.* 0.75) for the 3-Year (*resp.* 10-Year) maturity confirm that the estimated short-term and long-term liquidity premia are directly related to the liquidity costs and risks of green bond investors.

[INSERT FIGURE 3 HERE]

V. Drivers of liquidity premia

5.1. Joint dynamics of liquidity premia and economic determinants

Fig. 1 suggests that the evolutions of the short-term, long-term, and shape factors over the period 2015-2019 are correlated suggesting the presence of spillover effects. Assuming that liquidity shocks are larger for CGBs, spillover effects may occur at the short end of the liquidity premium and carry over to longer maturities. In order to test this possibility, we estimate the following unrestricted VAR-model denoted **Model 1** with 2 lags as in Kempf et al. (2012) :

$$\left[\begin{array}{l} \mathbf{Short}_t = \alpha_0^{so} + \sum_{i=1}^2 (\alpha_{i,so}^{so} \mathbf{Short}_{t-i} + \alpha_{i,sa}^{so} \mathbf{Shape}_{t-i} + \alpha_{i,l}^{so} \mathbf{Long}_{t-i}) + \varepsilon_t^{so} \\ \mathbf{Shape}_t = \alpha_0^{sa} + \sum_{i=1}^2 (\alpha_{i,so}^{sa} \mathbf{Short}_{t-i} + \alpha_{i,sa}^{sa} \mathbf{Shape}_{t-i} + \alpha_{i,l}^{sa} \mathbf{Long}_{t-i}) + \varepsilon_t^{sa} \\ \mathbf{Long}_t = \alpha_0^l + \sum_{i=1}^2 (\alpha_{i,so}^l \mathbf{Short}_{t-i} + \alpha_{i,sa}^l \mathbf{Shape}_{t-i} + \alpha_{i,l}^l \mathbf{Long}_{t-i}) + \varepsilon_t^l \end{array} \right. \quad \begin{array}{l} (4a) \\ (4b) \\ (4c) \end{array}$$

Where: The superscripts so, sa, and l represent the short, shape, and long factors, respectively.

$\varepsilon_t^{so}, \varepsilon_t^{sa}, \varepsilon_t^l$ denote the error terms.

We opt for a VAR (2) model in levels to capture level relations between the short, shape and long factors that determine the term structure of the liquidity premium.⁷ The column Model 1 of Table 6 reports the values of the Model 1 parameter estimates and their significance.

Overall, we shed light on a dynamic interaction between the short and long factors. Past short-term (*resp.* long-term) factors and long-term (*resp.* short-term) factors with one month lag determine the short-term (*resp.* long-term) liquidity premium. This first result indicates that the short end interacts with the long end of the term structure of the liquidity premium. A second important result is related to the positive coefficients of the shape factor with one lag on both short-term and long-term liquidity premia, which are significant at 5% level.

Taken together, these two results corroborate our conjecture about the existence of moderate spillover effects operating from the short end of the term structure to its long end via the shape factor as in Goyenko et al. (2011) consistent with **H₃**.

Hitherto, the interactions between the three premia help explain the dynamics of the liquidity premium but it is likely to be insufficient. Indeed, the literature has provided evidence of the impact of economic factors on the liquidity premium (e.g., Kempf et al., 2012). Unlike Broadstock and Cheng (2019), we employ the Economic policy uncertainty index (*EcoPoluncy*) measured for Europe to control for the uncertainty about economic policy.⁸ Furthermore, we select the monthly values of the Composite Leading Indicator (CLI) from OECD to account for the influence of the economic outlook in Europe (Kempf et al., 2012).

In addition, we consider the influence of the U.S. green bond market, which can be viewed by some investors as a substitute for the European green bond market. In fact, Zerbib (2019) and Tolliver et al. (2020) argue that tax exemptions attached to U.S. Municipal green bonds may attract long-term investors who seek to be hedged against inflation and illiquidity risks. Therefore, we construct a variable *USGreenMktPerf*, which captures the monthly variations of the S&P U.S. Municipal Green Bond index to account for the influence of the U.S. Municipal Green Bond market performance on investor strategies.

Finally, we use a set of control variables that may affect liquidity premia. A first set of variable controls for a changing liquidity difference between the CGB and the SSAGB segments. In this regard, we calculate the ratio volume issued in the CGB class divided by the total volume issued over the previous 6 months. As a result, we obtain *Ratio^{short}*, *Ratio^{medium}*, and *Ratio^{long}*, which capture the short end (bond maturity <5 years), the middle range (5 to 10

⁷ To account for a potential influence of past (monthly) changes in the liquidity premia, a lag length of two is selected. This value is suggested by the calculated Akaike (AIC) and Schwartz (SIC) information criteria.

⁸ Monthly data are available online: https://www.policyuncertainty.com/europe_monthly.html

years) and the long end of the term structure (>10 years) respectively. A second set of variables account for credit risk variations on the European corporate bond market. For that purpose, we consider the *Credit spread* variable. Unlike Kempf et al. (2012), we take end-of-month values for the 1-year bond maturity ($Credit^{short}$) at the short end. For the middle segment of the term structure, we take end-of-month values for 5-year maturity ($Credit^{medium}$) as well for the long end since no values are available for 10 years ($Credit^{long} = Credit^{medium}$).

To investigate the influence of the above-mentioned variables on the three liquidity premia, we incorporate them in the previous Model 1 in a VARX-model (denoted Model 2). Specifically, in this Model 2, we regress the three factors of the liquidity premium with 2 lags on the set of exogenous (economic and control) variables above presented such that:

$$\left[\begin{array}{l} \mathbf{Short}_t = \delta_0^{so} + \delta_1^{so} EcoOutlook_t + \delta_2^{so} EcoPoluncy_t + \delta_3^{so} USMunGreenPerf_t \\ \quad + \delta_4^{so} Credit_t^{short} + \delta_5^{so} Ratio_t^{short} + \sum_{i=1}^2 (\alpha_{i,so}^{so} Short_{t-i} + \alpha_{i,sa}^{so} Shape_{t-i} + \alpha_{i,sl}^{so} Long_{t-i}) + \varepsilon_t^{so} \\ \mathbf{Shape}_t = \delta_0^{sa} + \delta_1^{sa} EcoOutlook_t + \delta_2^{sa} EcoPoluncy_t + \delta_3^{sa} USMunGreenPerf_t \\ \quad + \delta_4^{sa} Credit_t^{medium} + \delta_5^{sa} Ratio_t^{medium} + \sum_{i=1}^2 (\alpha_{i,so}^{sa} Short_{t-i} + \alpha_{i,sa}^{sa} Shape_{t-i} + \alpha_{i,sl}^{sa} Long_{t-i}) + \varepsilon_t^{sa} \\ \mathbf{Long}_t = \delta_0^l + \delta_1^l EcoOutlook_t + \delta_2^l EcoPoluncy_t + \delta_3^l USMunGreenPerf_t \\ \quad + \delta_4^l Credit_t^{long} + \delta_5^l Ratio_t^{long} + \sum_{i=1}^2 (\alpha_{i,so}^l Short_{t-i} + \alpha_{i,sa}^l Shape_{t-i} + \alpha_{i,sl}^l Long_{t-i}) + \varepsilon_t^l \end{array} \right. \quad \begin{array}{l} (5a) \\ (5b) \\ (5c) \end{array}$$

Where: so, sa, and sl denote the short, shape, and long factors, respectively. $\varepsilon_t^{so}, \varepsilon_t^{sa}, \varepsilon_t^l$ are the error terms. *EcoOutlook*, *EcoPoluncy*, *USMunGreenPerf.*, *Credit* and *Ratio* are defined above.

The column Model 2 of Table 6 reports the coefficients of Model 2 variables and their significance. As for Model 1, those of short and long factors with one lag related to the short and long-term liquidity premia are positive and significant at 5% level. Those of shape factors with one lag related to short and long-term liquidity premia are also positive and significant.

The coefficients of *Economic outlook* are statistically significant and negative for the columns Short and Long. A lower figure of CLI index indicating pessimistic economic expectations contributes to increase short-term and long-term liquidity premia. According to our estimates, if the CLI index decreases by one standard deviation, the short-term and long-term liquidity premia increase by 4 bps i.e., 25% of the average liquidity premium. Another economic determinant of the short-term and long-term liquidity premia is *Economic policy uncertainty*. Given that its coefficients are significantly positive, the higher the *Economic policy uncertainty* is, the greater the short-term and long-term liquidity premia are. By contrast, the coefficients of *Economic outlook* and *Economic policy uncertainty* are not significant for the medium-term liquidity premium so it is not influenced by economic factors.

Interestingly, the coefficient of *U.S. Mun. Green Perf.* is significant and positive only for the long end of the liquidity premium. This result is important in the context of the development of competing regional green bond markets. We interpret this important result in two respects. On the one hand, investors are likely to require greater liquidity premium and thus better yields for long dated green bonds when the concurrent U.S. Municipal green bond market performs well. On the other hand, the implementation of tax incentives can decrease the compensation for higher illiquidity risks when they buy and hold long dated green bonds (Zerbib, 2019).

Turning to the control variables, we observe that the *Ratio* of recently issued bonds is insignificant for all three liquidity premia. Possibly, most investors do not regularly re-assess their notion of liquidity differential between CGBs and SSAGBs. In this regard, differences in issuance amount may be reflected into the constants of regression. This is the case of the long end of the term structure, since the positive and significant constant coincides with a greater issuance amount of long dated SSAGBs compared to long-dated CGBs (see Table 2). Finally, *Credit* seems to have no effect on liquidity premia. As in Longstaff (2004) for the U.S government bond market or in Kempf et al. (2012) and Schuster-Homburg for the German government bond market, parameter estimates are negative and insignificant for all premia.

Overall, Table 6 results indicate that economic factors determine the term structure of the liquidity premium in accordance with **H₂**. They also suggest moderate spillover effects running from the short end to the long end of the liquidity premium and vice-versa consistent with **H₃**.

[INSERT TABLE 6 HERE]

5.2. Bond age, economic determinants and liquidity premia

Analysing the term structure of the liquidity premium is important for investors when they consider maturity-specific liquidity premia within dynamic trading strategies for single bonds. Yet, when we regressed green bond yields given Eq. (1), we obtained significantly negative coefficients for *Bond Age* across maturities as in Mahanti et al. (2008) (see Table A1). This aging effect may be studied through the well-documented on-the-run vs. off-the-run effects. In fact, when a bond is recently issued, it is on-the-run, enjoying higher liquidity than when it has been outstanding for a while i.e., off-the-run. De Jong and Driessen (2015) note that studies generally documented a small (*resp.* greater) liquidity premium for the on-the-run (*resp.* off-the-run) treasury and government bonds. We test the possibility of the on-the-run vs. off-the-run effects by running the VARX model (Model 2) given two sub-samples: this composed of the on-the-run bonds that are recently issued (issuance have occurred for 2 years maximum) or not yet matured and this of the remaining bonds namely the off-the-run bonds.

Table 7 shows the values of coefficients and their significance for the two sub-samples. Looking at coefficients of lagged short, medium and long factors, their magnitude and signs are similar to those displayed in Table 6 for the on-the-run green bonds. In this way, the on-the-run-effect combined with spillover effects influence the term structure of the liquidity premium consistent with **H3**. Instead, the coefficients of short-term (*resp.* long-term) factors with one lag are only significant on the short-term (*resp.* long-term) liquidity premium for the off-the-run green bonds indicating the absence of spillover effects. Probably, this result is due to greater liquidity premia for the off-the-run bonds (De Jong and Driessen, 2015), which preclude the use of dynamic green bond trading strategies in favour of buy and hold strategies.

As seen in Table 6, the coefficients of *Economic Outlook* (*resp.* *Economic policy uncertainty*) are negative (*resp.* positive) and significant at 5% level for the short-term and long-term liquidity premia for the on-the-run bonds. By contrast, we report positive and significant coefficients of *U.S. Mun. Green Perf.* for the long-term liquidity premia for the two subsamples as in Table 6. Finally, we find that the short-term and long-term liquidity premia of the off-the-run bonds are insensitive to economic determinants probably because they are greater (De Jong and Driessen, 2015) and less risky since those bonds are closer to maturity.

[INSERT TABLE 7 HERE]

5.3. Liquidity premium and clientele effect

We now test the influence of the liquidity clientele effect on the liquidity premium of the green bond market. This clientele effect was detected on the German government bond market by Schuster and Uhrig-Homburg (2015). Then, Gehde-Trapp et al. (2018) show that it has opposite influences on the term structure of the ask and bid liquidity premia. In its presence, the ask liquidity premium increases at the short end and flattens out for longer maturities while the bid liquidity premium decreases at the short end and flattens out for longer maturities.

Following Gehde-Trapp et al. (2018), we obtain the ask and bid liquidity premia through the NS parameters estimated for the CGB and SSAGB class given ask and bid yields:

$$\begin{cases} \text{Liquiditypremium}(T)_{\text{ask},t} = \text{Yield}_{\text{ask},t}^{\text{CGB}}(T) - \text{Yield}_{\text{ask},t}^{\text{SSAGB}}(T) \\ \text{Liquiditypremium}(T)_{\text{bid},t} = \text{Yield}_{\text{bid},t}^{\text{CGB}}(T) - \text{Yield}_{\text{bid},t}^{\text{SSAGB}}(T) \end{cases} \quad (6)$$

Applying the same VARX model expressed in Eqs. (5a), (5b), (5c) allows us to test the joint impact of the liquidity investor clientele effect and spillover effects all in taking the influence of economic factors and investor strategies into account.⁹ As in Table 6, Table 8 shows short and long coefficients with one lag significant at 5% level for Short and Long

⁹ As explained by Zerbib (2019), the ask yield is the return investors earn if they paid the ask price and held the green bond to maturity while the bid yield is the return they earn if they sold it before maturity at the bid price.

columns respectively. Their magnitudes are noticeably greater for *ask yields*, which implies that buy and hold investors are confronted to more pronounced spillover effects running from the short end to the long end of the ask liquidity premium and vice-versa.

The column *ask yields* also displays significant coefficients for *Economic policy uncertainty* and *Economic outlook* with signs identical to those found in Table 6 for short-, medium-, and long-term premia, which indicate that those economic factors determine the buy and hold strategies of green bond investors. Conversely, the coefficients of *Economic outlook* are not significant for the medium and long-term liquidity premia for *bid yields*. This latest result suggests that the liquidity risks of investors, who seek to sell bonds before maturity with a maturity greater than 5 years are insensitive to the variations of the economic environment.

Moreover, we note that the short, medium and long factors have a significant and positive relationship with their respective coefficients estimated with one lag considering *ask yields*. These coefficients have a decreasing magnitude from short to long factors, which induces the shape of the ask liquidity premium predicted by Gehde-Trapp et al. (2018) i.e., the premium rises in the short-term and tends to flatten out in the long-term. This result is due to the influence of a liquidity investor clientele effect in accordance with **H4**. This clientele effect produces a (maturity) segmentation among buy and hold investors (Beber et al., 2021) because it implies that low liquidity risk (*resp.* high-liquidity risk) investors prefer buying short-dated (*resp.* long-dated) green bonds. By contrast, we see from the *bid yields* column that short and long factors have a significant and positive relationship with their respective coefficient with one lag whereas Gehde-Trappe et al. (2018) predicts a negative relationship. This result involves the absence of a liquidity clientele effect on the bid side. Accordingly, there is no segmentation among investors who are eager to sell green bonds before maturity. However, it reveals the existence of spillover effects running from the short end to the long end of the bid liquidity premium in accordance with **H3**, but not vice-versa as in Gehde-Trappe et al. (2018).

Owing to the potential endogenous relation between credit risk and liquidity, the liquidity premium and the liquidity clientele effect may be viewed as an effect of investors' risk preferences. As highlighted by Chen et al. (2020), funding constraints of some long-term investors like insurers may lead them to hold less risky bond portfolios (i.e., deleveraging risk effect). If so, the effect of such a bond risk clientele may supersede the liquidity clientele effect. However, we note that all coefficients of *Credit* are insignificant in Table 8. This latest result

clearly indicates that the clientele effect on the ask liquidity premium is not due to the credit risk channel but is related to the liquidity channel consistent with Chen et al. (2020).¹⁰

[INSERT TABLE 8 HERE]

5.4. Robustness tests

In this paragraph, we conduct two series of robustness tests assuming that the dynamics of liquidity premia may be differently affected by illiquidity events. In particular, the influence of economic factors could vary across different liquidity regimes: normal vs. stress (Schuster and Uhrig-Homburg, 2015; Smimou and Khallouli, 2016). In this regard, we consider the possibility of one or more breaks in the liquidity premia estimated from the VARX Model 2.

Specifically, we apply the Bai and Perron (1998, 2003) suite of tests to detect the presence of structural breaks in time series and ascertain their location. We search for a maximum of five breaks, the largest number allowed by the Bai and Perron procedure as it is common practice (e.g., Smimou and Khallouli, 2016 ; Karkowska and Urjasz, 2021).¹¹ The *UDmax* and the *WDmax* test statistics reported in Table 9 indicate that the null hypothesis of zero breaks is rejected at 5% in favour of the alternative one of an unknown number of breaks given a maximum of five breaks.¹² Then, we use the Bai and Perron sequential $SupF_{\tau}(L+1|L)$ test to determine the optimal number and location of breaks. The $SupF_{\tau}(1|0)$ is the only significant test for the three liquidity premia indicating that there is one structural break in their times series.

This break is identified in November 2017, where a spate of green bond issues propelled the world green bond market through the 100 billion euros barrier for the first time. This month notably saw two new and six repeat corporates come to the LGX market anticipating the One Planet Summit of Paris that took place in December 2017. This considerable flow of green bond issuances on LGX have implications on liquidity costs and illiquidity risks. In fact, we can see from Fig. 1 that the difference in bid-ask spreads between CGBs and SSAGBs decreased by 2 bps between November 2017 and January 2018 for the 2-Year and the 10-Year maturities. In the same period, the short-term and long-term liquidity premia decreased by almost 10%.

[INSERT TABLE 9 HERE]

We thus estimate the VARX Model 2 shown in Eqs. (5a), (5b), (5c) across two subperiods: January 2015-November 2017 and December 2017-December 2019.

¹⁰ We also run the VARX Model 2 given the two rating categories previously mentioned (AAA and AA or -) to control for bond-specific credit risks and obtained similar results.

¹¹ The trimming region provides the maximum possible number of breaks (=5 here) and minimum regime size.

¹² The *UDmax* statistic is the maximum value of the $SupF_{\tau}(L)$ F-statistic, while the *WDmax* statistic weights the individual statistics so as to equalize the p-values across values of $m = 5$ breaks.

Table 10 reports the values of estimated coefficients for the two periods. For the first period (Jan. 2015- Nov. 2017), the short and long coefficients with one lag are significant at 5% level for the short-term and long-term liquidity premia. This result implies spillover effects from the short end to the long end of the liquidity premium and vice versa as in Table 6 consistent with **H3**. The coefficients of *Economic Outlook*, *Economic Policy Uncertainty*, *U.S. Mun. Green Perf.* are also similar to those found in Table 6 in terms of sign and significance.

Moving to the second period (Dec 2017-Dec 2019), two interesting insights emerge. The first insight shows that the Long coefficient with one lag displayed in the column Short remains positive but becomes insignificant, suggesting the absence of spillover effects from the long end to the short end of the liquidity premium in contrast to Gehde-Trappe et al. (2018). The second insight relates to the insignificant coefficient of the *Economic Policy Uncertainty* variable that captures long-term liquidation risks of investors (Kempf et al., 2012). In the view of these two insights, the spillover effects and long-term liquidation risks of investors are unlikely to impact the long-term liquidity premium from 2018 consistent with Kempf et al. (2012). By contrast, the coefficients for *Economic Outlook* and *U.S. Mun. Green Perf.* variables are comparable to those found for the first period both in terms of sign and magnitude.

Overall, the results of Table 10 are qualitatively similar to those of Table 6 regarding the influence of economic factors confirming **H2**. This underscores the relevance of our VARX model to study the term structure of the liquidity premium of the green bond market.

[INSERT TABLE 10 HERE]

Our second robustness test is based on the estimation of the VARX Model 2 for the two subperiods by using ask and bid yields rather than by using close yields as in Table 10.

As seen in Table 8, Table 11 reports significant short and long coefficients with one lag at 5% level for the short and long liquidity premia whatever the period and the side (ask or bid). Since the magnitudes of those coefficients are higher for the period Nov. 2017 – Nov. 2019, investors face more pronounced spillover effects running from the short end to the long end of the liquidity premium and vice-versa in this period (Gehde et Trappe et al., 2018).

From Panel A, we observe that the short, medium and long factors have a significant and positive relationship with their respective coefficients estimated with one lag. As in Table 8, these lagged coefficients have a decreasing magnitude from short to long factors for the two periods, which bears out the prediction of Gehde-Trapp et al. (2018) related to the shape of the ask liquidity premium. This result also confirms the impact of a liquidity investor clientele effect for the two periods consistent with **H4**. This clientele effect induces a (maturity) segmentation among buy and hold investors (Beber et al., 2021), which leads low liquidity risk

(*resp.* high-liquidity risk) investors to prefer buying short-dated (*resp.* long-dated) green bonds. Finally, this clientele effect is related to the liquidity channel but not to credit risks in line with Table 8 results and Chen et al. (2020) given that all *Credit* coefficients are not significant.

From Panel B, we note that the coefficients of *Economic outlook* are always insignificant whatever the period selected. This suggests that the liquidity risks of investors, who seek to sell bonds before maturity are insensitive to the changes in economic prospects. Instead, both the coefficients and signs of *Economic Policy Uncertainty*, *U.S. Mun. Green Perf.* as well as these of short, shape and long factors with one lag are consistent with those found in Table 8.

Taken together, the results in Table 11 confirm and complement those of Table 8. They notably reveal that spillover effects but also the liquidity clientele effect are more pronounced from 2018, which involves a greater segmentation among buy and hold investors from 2018.

[INSERT TABLE 11 HERE]

VI. Conclusion

The green bond market provides a unique test bed for liquidity studies: Essentially highly rated bonds with green credentials that differ with respect to their liquidity, which are traded across all bond maturities. Our paper takes full advantage of those peculiarities to estimate the size of the liquidity premium and study its term structure. For that purpose, we compare the spread between yields of the more liquid SSAGB segment and the less-liquid CGB segments across different maturities. This spread reflects the (il)liquidity premium of the green bond premium across maturity, i.e., the term structure of liquidity premia.

Our empirical analysis of this liquidity premium delivers novel findings. First, our estimates of short-, medium- and long-term premia are close to those found on the German government bond market (Schuster and Uhrig-Homburg, 2015) respectively. However, the average difference between the long-term premium (16 bps) and the short-term one (13 bps) is lower than this obtained by Dick-Nielsen et al. (2012) for the corporate bond market. Second, we show that the short-term and long-term liquidity premia are sensitive to economic factors as in Kempf et al. (2012) but not to credit risks consistent with Chen et al. (2020). Economic outlook and uncertainty on economic policy gain importance for bonds with short and long maturities while the long-term liquidity premium is affected by tax-exemptions on U.S. green bonds. Third, we document that spillover effects operate from the short end to the long end of the liquidity premium and vice-versa (considering close or ask yields) and are more pronounced from 2018. With the on-the-run effect, they contribute to the U-shape of the term structure. Fourth, we find a liquidity clientele effect on the ask side, which impacts the term structure of

the ask liquidity premium. This clientele effect produces a (maturity) segmentation among buy and hold investors (Beber et al., 2021) since it leads low liquidity risk (*resp.* high liquidity risk) investors prefer buying short-dated (*resp.* long-dated) green bonds. However, there is no such a clientele effect on the bid side, which is indicative on the difficulties for short-term or long-term investors to sell green bonds before maturity in order to rebalance their portfolios or implement bond hedging strategies. In this respect, another implication of the liquidity clientele effect, which is that green bond ETFs may enhance the liquidity of the (underlying) green bonds through their direct investment or, as theory foresee, by cutting limits to arbitrage, is unlikely in contrast to the case of the U.S. corporate bond market (Dannhauser, 2017).

Our findings have both managerial and policy implications. From a managerial perspective, our findings suggest that issuers may select a liquidity investor clientele: high-risk (*resp.* low-risk) investors when issuing short-term (*resp.* long-term) bonds. This choice of an investor clientele will then impact their overall borrowing costs. Moreover, the investors' preference for buy and hold strategies may be a positive feature for issuers because it is unlikely that, during stress periods, investors will rush to sell their holdings, providing stability to the green bond market and preventing from large increase of liquidity premia during these periods as it was observed in the government bond market (Schuster and Uhrig-Homburg, 2015).

From a policy-making perspective, our results confirm the importance of the term structure of the liquidity premium for the functioning of secondary bond markets. Central bank interventions such as the purchases of green bonds by the European Central Bank (ECB hereafter) across different maturities may reduce liquidity risks so liquidity premia and yields, which could in turn stimulate the issuance of green bonds (Fender et al., 2019). Quite importantly, the term structure of our liquidity premium may be a relevant tool for ECB to avoid destabilising one of the two segments of the green bond market (CGBs and SSAGBs) for the benefit of the other segment by their purchases.

Two avenues for further research may be identified. A potential limitation of our study is that the liquidity differential between SSAGBs and CGBs may be partly explained by a residual credit risk premium assuming that SSAGBs are backed by supranational institutions and governments, which are very unlikely to default. Therefore, a possible extension of our work may be to decompose the yield premium into credit and liquidity premia. Another direction would be to assess the impact of the EU Taxonomy recently voted and the related EU green bond standard that allow for greater transparency on the use of green bond proceeds (Lebelle et al., 2021) on the liquidity premium and on investor strategies.

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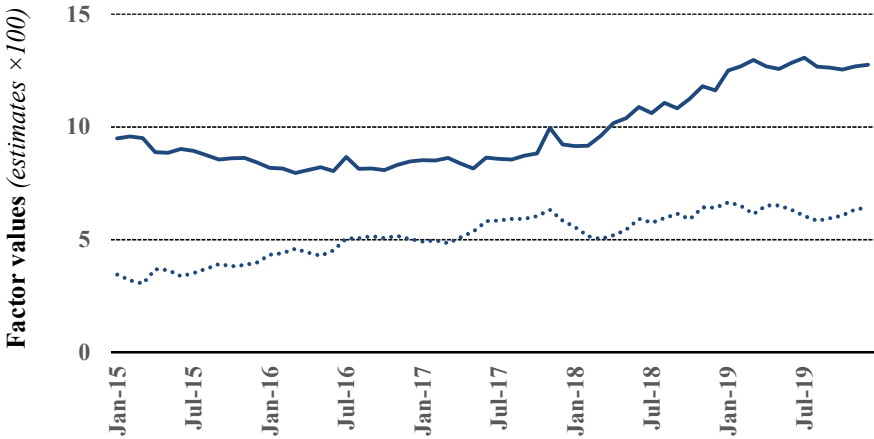
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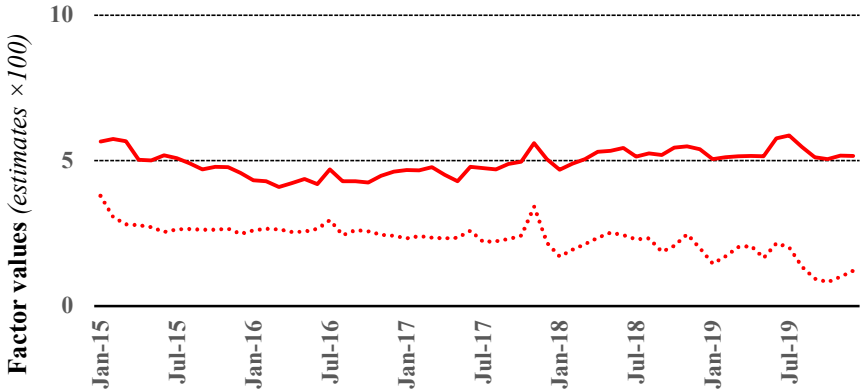
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VIII. Tables and Figures

Part A : Long-term factors



Part B : Short-term factors



Part C : Shape factors

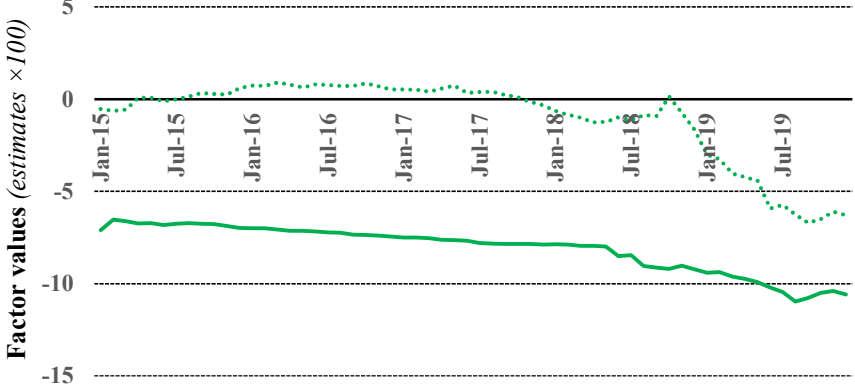


Fig. 1. Term structure factors. This figure presents the evolution of the term structure factors estimated over a 5-year period (January 2015-December 2019). The solid lines depict the factors for the Corporate (CGB) segment and the dashed lines the factors for the segment of Supranationals, Sovereign and Agencies (SSAGB) segment. Part A provides the estimates of the long-term factors ($\beta_{0,t}$) ; Parts B and C the estimates of the short-term factors ($\beta_{1,t}$) and the shape factors ($\beta_{2,t}$) respectively.

Estimations are run by minimizing the sum of squared yield differences according to Eq. (1) for end-of-month yields with maturities of 2, 3, 4, 5, 6, 7, 8, 9, 10, 12, 13, 15 and 20 years for each segment.

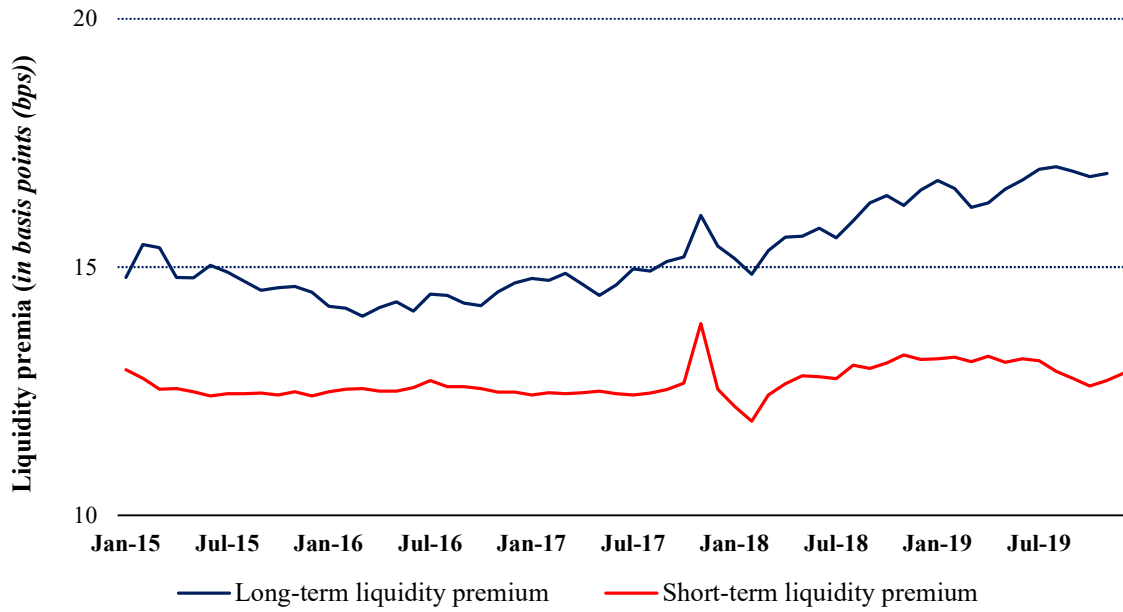


Fig 2. Short-term and long-term liquidity premia (fitted curves). This figure shows the evolution of the short-term and the long-term liquidity premia over the period: January 2015-December 2019 (60 monthly observations). The red line outlines the short-term liquidity premium, which is the difference between the short-term factors of the Corporate Green Bond (CGB) segment and those of the Sovereign Green Bond (SSAGB) segment i.e., $\beta_{0,t}^{CGB} - \beta_{0,t}^{SSAGB} + \beta_{1,t}^{CGB} - \beta_{1,t}^{SSAGB}$. The blue line depicts the long-term liquidity premium that is the difference between the long-term factors of the CGB and SSAGB segment respectively i.e., $\beta_{0,t}^{CGB} - \beta_{0,t}^{SSAGB}$.

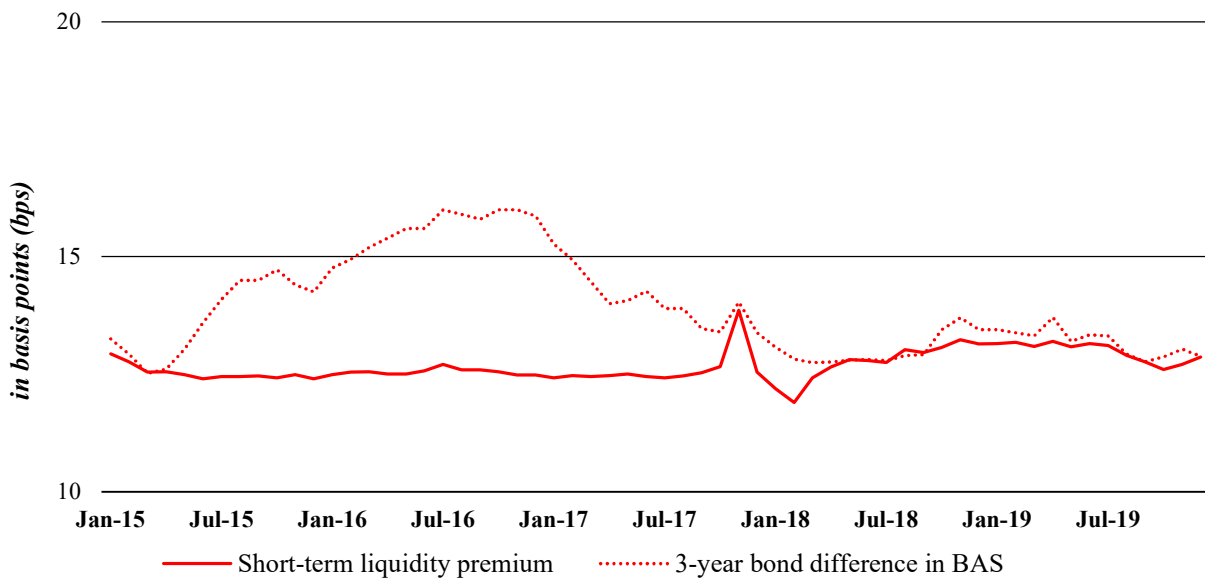
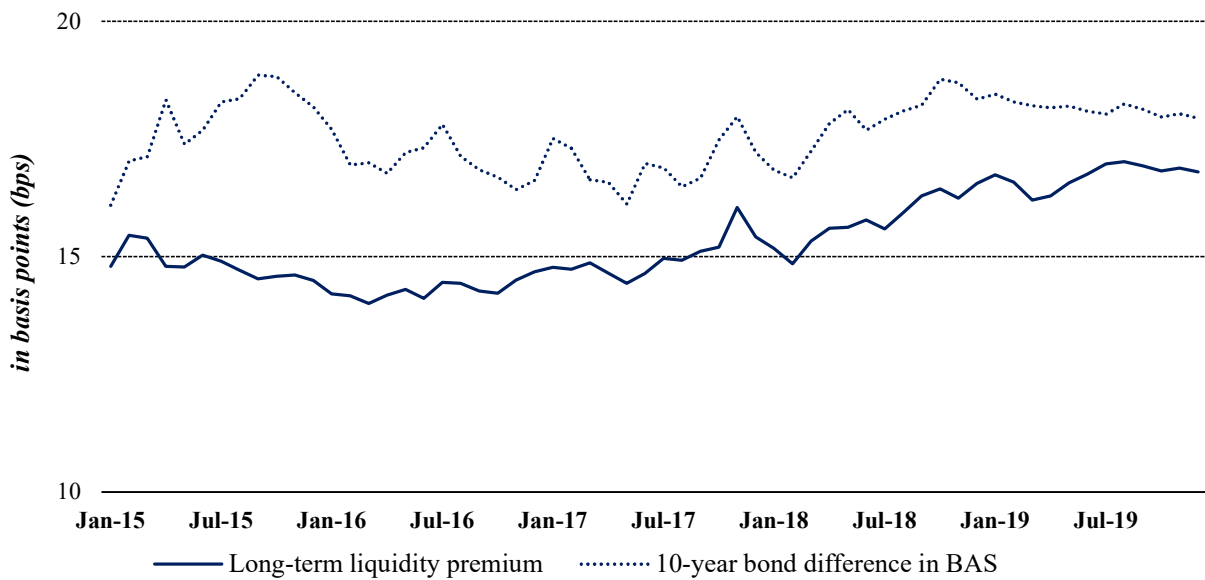


Fig. 3. Liquidity premiums (fitted curves) and bid-ask spreads. This figure plots the evolution of liquidity premiums (solid lines) together with the absolute bid-ask spread differences between CGBs and SSAGBs (dotted lines) over the period January 2015-December 2019 (60 monthly observations). The upper graph depicts the evolution of the difference in absolute bid-ask spreads (difference in BAS) and this of liquidity premium of bonds with a maturity of ten years. The lower graph presents the evolution of the difference in absolute bid-ask spreads and this of liquidity premium of bonds with a maturity of 3 years.

Table 1. Green Bond primary market and European exchanges activity (2018-2019)

Name	Type of section	2018 Issuance (in USD Million)	2019 Issuance (in USD Million)	Average market share (in %)
LuxSE/LGX*	Green bonds	25	28	25.7%
LSEG (LSE & Borsa Italiana)	Green and Social bonds	18.2	18.25	17.7%
German Exchanges	Green bonds	10.24	15.2	12.3%
HFEX	Green and Sustainable bonds	5.2	6.3	5.6%
Euronext**	Green and Sustainable bonds	14.66	27.5	20.4%
Swiss Stock Exchange	Green bonds	4.52	3.9	4.1%
SGX	Green and Social bonds	3.67	6.8	5.1%
Nasdaq Nordic	Green and Sustainable bonds	3.64	3.75	3.6%
Others***		5.34	6.03	5.5%
TOTAL		90.47	115.73	100%

Notes: The average market share is the proportion of bond issuance amount for a given exchange among the full amount issued by exchanges between 2018 and 2019.

* The green bond listing of LGX (Luxemburg Green Exchange) has been active since September 2016.

** Euronext exchanges include Paris (9.3%), Dublin (6.1%), Brussels (2.2%) and Amsterdam (2.8%) platforms.

*** Other German stock exchanges are the following: Frankfurt (5.1%), Berlin (3.5%), Stuttgart (2%), Munich (1%), Dusseldorf (0.7%) and are owned by the Deutsche Boerse Group.

Table 2. Descriptive statistics of the green bond sample*Panel A: Summary statistics of liquidity and credit variables*

	SSAGBs	CGBs	Total
Number of bonds	93	91	184
Issue Size (<i>in million EUR</i>)	837.785	605.955	723.494
Maturity (<i>in years</i>)	7.34	6.95	7.10
Coupon (<i>in %</i>)	3.14	2.78	2.97
Rating	1.20	1.70	1.46
Closing yields (<i>in %</i>)	1.22	1.32	1.25
Ask yields (<i>in %</i>)	1.19	1.24	1.21
Bid yields (<i>in %</i>)	1.44	1.52	1.48
Bid-ask spread (BAS) (<i>in bps</i>)	31.98	39.71	35.82
Bid-ask spread (BAS) (<i>in %</i>)	0.292	0.405	0.350
ZTD (<i>in %</i>)	3.23	7.15	5.1
Roll measure	0.187	0.425	0.317

Panel B: Distribution statistics for liquidity proxies

	BAS (<i>in %</i>)	Roll	ZTD (<i>in %</i>)
95th	0.931	0.993	13.66
75th	0.540	0.666	8.6
Median	0.322	0.392	5.01
25th	0.151	0.091	2.69
5th	0.093	0.028	1.76

Notes: Monthly Bid-Ask spread (BAS) are calculated using quote data provided by Bloomberg, resulting in 60 observation months over the five-year period January 1, 2015 to December 30, 2019. The average closing yield is computed with closing prices while ask yields and bid yields are calculated with ask and bid prices.

The average rating is assessed on the basis of the Moody's rating obtained for the green bond issued (or on the basis of the S&P rating if the Moody's rating is not available) given the following scale : 1 (*Prime* – rating : AAA), 2 (*High Grade* – rating: Aa1, Aa2 or Aa3), 3 (*Upper Medium Grade* – rating: A1, A2 or A3),..., 9 (*Default imminent* – rating : Caa3, Ca or C), and 10 (*In default* – rating D).

ZTD is the proportion of days with zero returns within a month. The Roll measure is based on the covariance between subsequent (daily) returns.

Table 3. Liquidity regressions*Panel A: With bond (close) yields*

	Full sample	AAA	AA	A and BBB
BAS	0.0821** (4.15)	0.0902** (5.12)	0.0754** (3.96)	0.0786** (4.87)
ZTD	0.0006* (2.14)	0.0007* (4.20)	0.001** (2.31)	0.0008* (2.23)
Roll	0.0427** (3.18)	0.0412** (2.78)	0.0391* (2.02)	0.0435** (7.22)

Panel B: With ask (close) yields

	Full sample	AAA	AA	A and BBB
BAS	0.0871** (3.88)	0.0917** (4.46)	0.077** (2.98)	0.0802** (3.70)
ZTD	0.0003 (1.38)	0.0006 (2.08)	0.002 (0.90)	0.0003 (1.76)
Roll	0.0442** (2.92)	0.0436** (2.61)	0.0348* (1.99)	0.0473** (6.13)

Panel C: With bid (close) yields

	Full sample	AAA	AA	A and BBB
BAS	0.0818** (3.55)	0.088** (4.37)	0.0732* (2.05)	0.0776** (3.65)
ZTD	0.0005 (1.17)	0.0004 (1.21)	-0.0002 (0.56)	0.0006 (1.30)
Roll	0.0383** (3.02)	0.0405** (2.63)	0.0268 (1.84)	0.0340** (5.76)

Note: For three rating classes (AAA, AA, A and BBB) and for the full sample, we run the regression model set in Eq. (1) by considering BAS_{it} , ZTD_{it} and $Roll_{it}$ as liquidity variables, respectively. 12 regressions are therefore run (namely 3 liquidity variables \times 4 rating classes).

This table shows for all regressions the coefficient estimates and t-statistics in parenthesis for the liquidity variable. Standard errors are corrected for time series effects, firm fixed effects, and heteroscedasticity.

Statistical significance of coefficients at 5% (*resp.* 1%) level is marked * (*resp.* **).

Table 4. Size of the liquidity premium.*Panel A: Liquidity premium estimated using the method of Dick-Nielsen et al. (2012)*

	Average (in bps)	Liquidity premium (in bps)			Number of observations		
		2-5y	5-10y	10-20y	2-5y	5-10y	10-20y
Full sample	17.9	17.0 (16.1 ; 17.9)	16.5 (15.4 ; 17.6)	20.5 (19.5; 21.5)	81	48	55
AAA	17.6	16.8 (16.2 ; 17.5)	16.3 (15.5 ; 17.1)	20.3 (19.3 ; 21.3)	60	28	33
AA and -	18.2	17.1 (16.0 ; 18.2)	16.7 (15.4 ; 18.0)	20.7 (19.7; 21.7)	21	20	22

Panel B: Liquidity premium estimated using the difference between CGB and SSAGB samples

	Average (in bps)	Liquidity premium (in bps)			Number of observations		
		2-5y	5-10y	10-20y	2-5y	5-10y	10-20y
Full sample	13.9	13.3 (12.7 ; 13.9)	12.7 (11.6 ; 13.8)	15.8 (14.9; 16.7)	81	48	55
AAA	13.7	13.2 (12.7 ; 13.7)	12.5 (11.5 ; 13.6)	15.6 (14.8; 16.4)	60	28	33
AA and -	14.2	13.4 (12.6 ; 14.2)	12.8 (11.7; 14.0)	15.9 (14.9; 16.9)	21	20	22

Notes: For every rating R we run the pooled regression: $\text{Yield}_{it} = \alpha R + \beta \text{BAS}_{it} + \text{controls}_{it} + \epsilon_{it}$ where i refers to bond, t to time, and BAS_{it} is the bid-ask spread measure (denoted BAS).

Panel A displays the values of the liquidity premiums estimated according to the method of Dick-Nielsen (2012). We sort increasingly all values of BAS_{it} and find the median value BAS_{50} and the 5% value BAS_5 within each rating and maturity bucket (2-5y, 5-10y, and 10-20y). The liquidity component is defined as $\beta (\text{BAS}_{50} - \text{BAS}_5)$.

Panel B reports the values of liquidity premiums estimated using the difference between the median BAS of CGB sample and that of the SSAGB sample within each rating and maturity bucket (2-5y, 5-10y, and 10-20y).

The liquidity component in the bucket is defined as $\beta (\text{BAS}_{50(\text{CGB})} - \text{BAS}_{50(\text{SSAGB})})$.

Both Panels A and B show for all buckets the liquidity premium component with standard errors in parenthesis. Confidence bands are found by a wild cluster bootstrap.

Table 5. Liquidity component in fraction of the bond yield.*Panel A: Liquidity component estimated using the method of Dick-Nielsen et al. (2012)*

	Fraction of the yield (<i>in %</i>)			Number of observations		
	2-5y	5-10y	10-20y	2-5y	5-10y	10-20y
Full sample	13 (12;14)	12 (10;14)	16 (14;18)	81	48	55
AAA	13 (12;14)	12 (10;14)	15.5 (14;17)	60	28	33
AA and -	13.5 (12;15)	12 (10;14)	16 (14;18)	21	20	22

Panel B: Liquidity component estimated using the difference between CGBs and SSAGBs

	Fraction of the yield (<i>in %</i>)			Number of observations		
	2-5y	5-10y	10-20y	2-5y	5-10y	10-20y
Full sample	10 (9;11)	9 (7;11)	12 (11;13)	81	48	55
AAA	10 (9;11)	9 (7;11)	12 (11;13)	60	28	33
AA and -	10.5 (8;13)	9 (6;12)	12.5 (10;14)	21	20	22

Notes: For every rating R we run the pooled regression: $\text{Yield}_{it} = \alpha_R + \beta \text{BAS}_{it} + \text{controls}_{it} + \varepsilon_{it}$ where i refers to bond, t to time, and BAS_{it} is the bid-ask spread measure (denoted BAS).

Panel A displays the values of the liquidity components estimated according to the method of Dick-Nielsen (2012) (see Table 4). For each bond we define the liquidity fraction of the bond yield as $\beta^R (\text{BAS}_{i,t} - \text{BAS}_5) / \text{Yield}_{i,t}$.

The estimated fractions in bold are the median fraction.

Panel B reports the values of liquidity components estimated using the BAS difference between the CGB and SSAGB samples. For each bond, we compute the liquidity fraction of the yield as $\beta^R \text{BAS}_{i,t} / \text{Yield}_{i,t}$. The estimated fractions in bold are the difference between the median of CGB sample and that of SSAGB sample.

Confidence bands in parentheses are found by a wild cluster bootstrap.

Table 6. Liquidity premia: economic determinants and joint dynamics

	Model 1			Model 2		
	Short	Shape	Long	Short	Shape	Long
<i>Constant</i>	-3.309 (2.543)	1.993 (3.791)	1.270** (0.476)	76.17 (45.08)	-134.42* (64.89)	9.59 (8.76)
<i>Short (t-1)</i>	2.265** (0.820)	-2.690* (1.223)	0.335* (0.154)	1.79** (0.64)	-1.93 (1.13)	0.325* (0.152)
<i>Short (t-2)</i>	-1.397* (0.650)	2.411* (1.119)	-0.167 (0.141)	0.312 (0.763)	-0.340 (1.190)	0.099 (0.161)
<i>Shape (t-1)</i>	0.761* (0.545)	-0.639 (0.812)	0.209* (0.102)	0.853* (0.385)	-0.610 (0.729)	0.217* (0.098)
<i>Shape (t-2)</i>	-0.627 (0.496)	1.418 (0.739)	-0.094 (0.092)	0.231 (0.497)	0.319 (0.772)	0.073 (0.104)
<i>Long (t-1)</i>	1.315* (0.676)	-2.621* (1.231)	1.228** (0.155)	1.195* (0.531)	-2.191* (1.182)	1.094** (0.173)
<i>Long (t-2)</i>	-0.917 (0.687)	1.686 (1.025)	-0.136 (0.129)	0.251 (0.672)	-0.143 (1.063)	0.104 (0.144)
<i>Ratio</i>				-2.847 (3.407)	2.121 (4.904)	0.198 (0.446)
<i>Credit</i>				0.041 (0.0495)	-0.052 (0.045)	-0.002 (0.002)
<i>Economic Outlook</i>				-0.082** (0.038)	1.210 (0.887)	-0.089* (0.040)
<i>Economic policy Uncertainty</i>				0.0024* (0.001)	-0.002 (0.002)	0.002** (0.0009)
<i>U.S. Mun. Green Perf.</i>				0.097 (0.068)	-0.136 (0.094)	0.015** (0.004)
R²	0.94	0.95	0.91	0.97	0.97	0.93

Notes: The column **Model 1** presents the results for the VAR (2) Model 1 shown in Eqs. (3a), (3b) and (3c) that is estimated over the period January 2015–December 2019 (60 observations).

The column **Model 2** displays the results for the VARX Model 2 outlined in Eqs. (4a), (4b) and (4c) that is estimated over the period January 2015–December 2019 (60 observations).

Standard errors of the coefficients are expressed in parentheses and they are based on the Newey and West's (1987) covariance matrix estimator with 6 lags.

* (*resp.* **) means that the coefficient is significant at 5% (*resp.* 1%) level.

Table 7. Liquidity premia: bond age, economic determinants and joint dynamics

	On-the-run bonds			Off-the run-bonds		
	Short	Medium	Long	Short	Medium	Long
<i>Constant</i>	1.456** (0.480)	1.005 (1.453)	1.731* (0.786)	0.567* (0.238)	-0.935 (0.887)	0.18 (0.195)
<i>Short (t-1)</i>	0.402** (0.160)	-1.434 (4.877)	0.771* (0.362)	0.698** (0.234)	-2.887 (2.322)	0.289 (0.201)
<i>Short (t-2)</i>	0.077 (0.182)	-0.944 (3.788)	-0.915 (1.970)	0.055 (0.229)	1.810 (1.970)	0.161 (0.325)
<i>Shape (t-1)</i>	0.085* (0.037)	-0.845 (0.592)	0.123* (0.052)	0.014 (0.02)	0.043 (0.234)	0.110* (0.092)
<i>Shape (t-2)</i>	-0.003 (0.005)	0.033 (0.169)	0.009 (0.065)	-0.005 (0.021)	0.022 (0.144)	-0.058 (0.103)
<i>Long (t-1)</i>	0.112* (0.044)	0.401 (0.823)	0.674** (0.244)	0.076 (0.045)	-0.182 (0.311)	0.566* (0.249)
<i>Long (t-2)</i>	0.016 (0.028)	-0.567 (0.612)	0.080 (0.177)	-0.007 (0.055)	0.093 (0.211)	0.110 (0.178)
<i>Ratio</i>	-0.013 (0.025)	-0.202 (0.245)	0.023 (0.066)	-0.009 (0.013)	-0.007 (0.019)	0.017 (0.058)
<i>Credit</i>	-0.021 (0.056)	0.133 (0.208)	-0.087 (0.066)	-0.009 (0.017)	0.386 (0.377)	-0.041 (0.123)
<i>Economic Outlook</i>	-0.675* (0.310)	-0.070 (0.101)	-0.423* (0.201)	-0.391* (0.142)	0.077 (0.134)	-0.162* (0.082)
<i>Economic Policy Uncertainty</i>	0.0022* (0.001)	-0.0001 (0.001)	0.0013* (0.0005)	0.0018 (0.002)	0.002 (0.016)	0.002 (0.003)
<i>U.S. Mun. Green Perf.</i>	0.135 (0.157)	-0.098 (0.245)	0.111* (0.046)	0.096 (0.082)	-0.654 (0.990)	0.154* (0.076)
R²	0.88	0.86	0.89	0.82	0.78	0.84

Notes: This table shows the results of the VAR Model 2 shown in Eqs. (5a), (5b), (5c) estimated over the period: January 2015–December 2019 (60 observations) for two subsamples.

The sample of the **on-the-run bonds** contain bonds that are recently issued (issuance have occurred for 1 year at least and 2 years maximum) and have not yet matured. The sample of the **off-the-run bonds** contain bonds that have either have a shorter time to maturity or have already matured.

The exogenous variables are defined as in Table 6.

Standard errors of the coefficients are in parentheses and are based on the Newey and West's (1987) covariance matrix estimator with 6 lags.

* (*resp.* **) means that the coefficient is significant at 5% (*resp.* 1%) level.

Table 8. Ask and bid liquidity premia: economic determinants and joint dynamics

	Ask yields			Bid yields		
	Short	Medium	Long	Short	Medium	Long
<i>Constant</i>	0.351* (0.166)	-1.012 (0.889)	-0.332 (0.601)	0.132** (0.045)	1.765 (1.551)	-0.489 (0.745)
<i>Short (t-1)</i>	1.267** (0.448)	2.019* (0.924)	0.718* (0.337)	0.423* (0.197)	-2.463 (5.121)	1.132 (1.34)
<i>Short (t-2)</i>	0.088 (0.432)	-0.476 (0.854)	0.043 (0.196)	0.122 (0.356)	-2.082 (4.237)	-1.076 (1.002)
<i>Shape (t-1)</i>	0.425 (0.320)	1.946* (1.020)	0.313* (0.157)	0.087 (0.190)	0.145 (0.208)	0.097 (0.161)
<i>Shape (t-2)</i>	0.013 (0.422)	0.413 (0.987)	-0.134 (0.881)	-0.034 (0.123)	-0.076 (0.223)	-0.09 (0.102)
<i>Long (t-1)</i>	0.686 (0.367)	1.227* (0.549)	0.802** (0.323)	0.032 (0.047)	-0.211 (0.454)	0.754** (0.301)
<i>Long (t-2)</i>	0.097 (0.658)	-0.233 (1.112)	-0.097 (0.201)	-0.050 (0.111)	0.047 (0.655)	-0.177 (0.256)
<i>Ratio</i>	-0.298 (0.755)	-3.442 (4.677)	-0.123 (0.255)	-0.008 (0.02)	0.178 (0.323)	0.033 (0.103)
<i>Credit</i>	-0.031 (0.021)	-0.012 (0.057)	-0.009 (0.087)	0.008 (0.096)	-0.002 (0.011)	0.010 (0.009)
<i>Economic Outlook</i>	-0.886** (0.302)	-0.655 (0.445)	-0.227* (0.093)	-0.243* (0.112)	-0.076 (0.064)	0.112 (0.123)
<i>Economic Policy Uncertainty</i>	0.0026** (0.001)	-0.001 (0.002)	0.0015* (0.001)	0.0019* (0.001)	0.002 (0.004)	0.0008 (0.0005)
<i>U.S. Mun. Green Perf.</i>	0.149 (0.171)	-0.145 (0.234)	0.198* (0.087)	0.120 (0.101)	0.153 (0.745)	0.096* (0.088)
R²	0.84	0.89	0.92	0.86	0.86	0.93

Notes: This table presents the results of the VARX Model 2 shown in (5a), (5b), (5c) estimated over the period: January 2015–December 2019 (60 observations) based on either on the ask yields or on the bid yields.

The exogenous variables are defined as in Table 5.

Standard errors of the coefficients are in parentheses and are based on the Newey and West's (1987) covariance matrix estimator with 6 lags.

* (resp. **) means that the coefficient is significant at 5% (resp. 1%) level.

Table 9. Estimated number of breaks and break dates for liquidity premia.

	NS(1987) estimates			Estimated from original data		
	Short	Shape	Long	3-Year	5-Year	15-Year
<i>UDmax</i>	19.93**	209.68**	99.07**	7.96	207.93**	37.80**
<i>WDmax_{5%}</i>	19.93**	209.68**	99.07**	9.96	207.93**	37.80**
<i>SupF_T(1 0)</i>	19.93**	209.68**	99.07**	9.96	207.93**	37.80**
<i>SupF_T(2 1)</i>	2.08	9.48	9.44	8.84	6.25	3.80**
<i>SupF_T(3 2)</i>	4.27	10.27	4.24	8.42	2.36	8.09**
Number of breaks	1	1	1	0	1	1
Break Date	2017:11	2017:11	2017:11		2017:11	2017:11

Note: This table reports the double maximum test of breaks in the means of three illiquidity premia estimated with the NS model (short, shape, long) or obtained from original data (3-Year, 5-Year and 15-Year premium). Two double maximum tests that test the null hypothesis of no structural break —*UDmax* and *WDmax*—are reported in the first two rows. Rows 3–5 report the sequential *SupF_T(L + 1/L)* *F*-statistic for *L*=1,2,3 to determine the number and location of structural breaks, The location are determined according to the Bai and Perron (1998) critical values, which are 8.58, 10.13 and 11.14 for *SupF_T(1|0)*, *SupF_T(2|1)*, *SupF_T(3|2)* respectively given a 15% trimming rate. The sample period is from January 2015 to December 2019 on a monthly basis (60 observations).

* (resp. **) means that the coefficient is significant at 5% (resp. 1%) level.

Table 10. VARX Model 2 estimated over two subperiods: January 2015–November 2017 vs. December 2017–December 2019

	Jan. 2015 – Nov. 2017			Dec. 2017 – Dec. 2019		
	Short	Medium	Long	Short	Medium	Long
<i>Constant</i>	-61.31 (32.53)	46.32 (45.86)	12.74 (6.49)	-125.52 (123.67)	-26.98 (296.7)	8.04 (40.81)
<i>Short (t-1)</i>	0.729** (0.221)	-1.149 (1.017)	0.223* (0.095)	1.588* (0.690)	-2.999 (2.903)	0.559* (0.261)
<i>Short (t-2)</i>	-0.097 (0.475)	-0.023 (0.670)	0.084 (0.143)	0.835 (1.945)	-1.382 (2.878)	0.120 (0.288)
<i>Shape (t-1)</i>	0.346* (0.170)	-0.735 (0.804)	0.162* (0.073)	0.828 (0.519)	-1.748 (1.807)	0.155* (0.069)
<i>Shape (t-2)</i>	-0.041 (0.339)	-0.005 (0.479)	0.045 (0.068)	0.536 (1.233)	-0.835 (1.826)	0.082 (0.309)
<i>Long (t-1)</i>	0.656* (0.321)	-1.775* (0.83)	0.299* (0.147)	0.578 (0.380)	-1.96 (1.27)	1.108* (0.423)
<i>Long (t-2)</i>	0.328 (0.397)	-0.347 (0.560)	0.087 (0.079)	1.386 (2.67)	-2.14 (4.02)	0.427 (0.675)
<i>Ratio</i>	-1.260 (2.251)	-0.151 (3.174)	0.109 (0.449)	5.992 (6.114)	-28.18 (33.53)	-1.070 (3.593)
<i>Credit</i>	0.011 (0.009)	-0.016 (0.013)	-0.001 (0.002)	0.080 (0.053)	-0.135 (0.078)	0.007 (0.013)
<i>Economic Outlook</i>	-0.661** (0.313)	-0.663* (0.311)	-0.091* (0.043)	-0.752* (0.302)	-0.645* (0.311)	-0.083* (0.042)
<i>Economic Policy Uncertainty</i>	0.002* (0.0007)	0.002* (0.0008)	0.0003* (0.0001)	0.002* (0.001)	0.006* (0.003)	0.001 (0.001)
<i>U.S. Mun. Green Perf.</i>	0.15 (0.211)	-0.050 (0.059)	0.015* (0.07)	0.303 (0.206)	-0.162 (0.103)	0.019* (0.009)
R²	0.89	0.86	0.83	0.91	0.93	0.93

Notes: This table shows the results of the VARX Model 2 that is estimated over two periods: January 2015–November 2017 (35 observations) and December 2017–December 2019 (25 observations).

Standard errors of the coefficients are in parentheses and are based on the Newey and West's (1987) covariance matrix estimator with 6 lags.

* (*resp.* **) means that the coefficient is significant at 5% (*resp.* 1%) level.

Table 11. VARX Model 2 estimated over two subperiods: January 2015–November 2017 vs. December 2017–December 2019

Panel A: Using ask yields

	Jan. 2015 – Nov. 2017			Dec. 2017 – Dec. 2019		
	Short	Medium	Long	Short	Medium	Long
<i>Constant</i>	8.201 (6.233)	-7.472 (5.88)	-5.672 (4.322)	1.351 (1.612)	-3.012 (2.703)	-2.932 (1.677)
<i>Short (t-1)</i>	1.105** (0.334)	1.087 (0.742)	0.652* (0.318)	1.365** (0.403)	2.55* (1.087)	0.767* (0.337)
<i>Short (t-2)</i>	0.066 (0.861)	-0.288 (0.943)	0.03 (0.234)	0.121 (0.675)	-0.534 (1.102)	0.087 (0.226)
<i>Shape (t-1)</i>	0.367 (0.281)	1.533* (0.612)	0.278* (0.124)	0.513 (0.411)	1.425* (0.638)	0.435* (0.191)
<i>Shape (t-2)</i>	0.009 (0.249)	0.234 (0.766)	-0.022 (0.187)	0.02 (0.071)	0.641 (1.788)	-0.223 (0.673)
<i>Long (t-1)</i>	0.577 (0.361)	1.012* (0.492)	0.283** (0.089)	0.832 (0.578)	1.58 (1.184)	0.934** (0.245)
<i>Long (t-2)</i>	-0.087 (0.335)	0.101 (1.003)	0.088 (0.341)	0.191 (0.779)	-0.208 (0.812)	0.134 (0.557)
<i>Ratio</i>	-0.323 (0.834)	-1.895 (1.451)	-0.187 (0.378)	-0.175 (0.434)	-2.081 (1.452)	-0.144 (0.228)
<i>Credit</i>	-0.033 (0.038)	-0.027 (0.102)	-0.035 (0.047)	-0.018 (0.045)	-0.01 (0.086)	-0.017 (0.033)
<i>Economic Outlook</i>	-0.637* (0.288)	0.101 (0.546)	-0.106* (0.049)	-0.907** (0.296)	-0.422 (0.566)	-0.362* (0.135)
<i>Economic Policy Uncertainty</i>	0.002** (0.001)	0.004 (0.004)	0.001 (0.001)	0.003* (0.001)	-0.005 (0.006)	0.002* (0.001)
<i>U.S. Mun. Green Perf.</i>	0.254 (0.222)	-0.364 (0.453)	0.097* (0.039)	0.455* (0.191)	-0.087 (0.24)	0.259* (0.121)
R²	0.80	0.85	0.90	0.84	0.87	0.87

Panel B: Using bid yields

	Jan. 2015 – Nov. 2017			Dec. 2017 – Dec. 2019		
	Short	Medium	Long	Short	Medium	Long
<i>Constant</i>	1.23* (0.597)	2.761 (1.897)	-1.761 (1.237)	0.989* (0.466)	2.133 (1.788)	0.778 (0.924)
<i>Short (t-1)</i>	0.288* (0.135)	-3.677 (4.234)	0.788 (1.796)	0.602** (0.215)	-2.003 (3.335)	0.872 (1.282)
<i>Short (t-2)</i>	0.176 (0.484)	-1.522 (2.671)	-0.78 (1.986)	0.08 (0.289)	1.113 (2.788)	-0.651 (1.791)
<i>Shape (t-1)</i>	0.055 (0.077)	0.087 (0.262)	0.04 (0.161)	0.133 (0.162)	0.267 (0.425)	0.201* (0.092)
<i>Shape (t-2)</i>	0.142 (0.287)	0.164 (0.319)	-0.199 (0.645)	-0.102 (0.226)	-0.244 (0.358)	-0.024 (0.187)
<i>Long (t-1)</i>	0.025 (0.036)	0.098 (0.335)	0.433* (0.208)	0.047 (0.039)	-0.033 (0.543)	0.94** (0.278)
<i>Long (t-2)</i>	0.228 (0.821)	0.172 (0.648)	-0.323 (0.441)	-0.064 (0.176)	0.098 (0.143)	-0.121 (0.286)
<i>Ratio</i>	0.087 (0.103)	-0.075 (0.16)	0.021 (0.088)	-0.023 (0.097)	0.249 (0.491)	0.138 (0.465)
<i>Credit</i>	0.004 (0.006)	0.003 (0.008)	0.008 (0.011)	0.012 (0.023)	-0.004 (0.01)	0.019 (0.022)
<i>Economic Outlook</i>	-0.332* (0.148)	0.141 (0.167)	0.23 (0.274)	-0.134 (0.216)	-0.088 (0.252)	0.073 (0.092)
<i>Economic Policy Uncertainty</i>	0.004* (0.002)	-0.001 (0.003)	0.002 (0.002)	0.003* (0.001)	0.004 (0.007)	0.002* (0.001)
<i>U.S. Mun. Green Perf.</i>	0.078 (0.092)	0.087 (0.231)	0.062* (0.027)	0.215 (0.188)	0.276 (0.368)	0.185* (0.117)
R²	0.78	0.80	0.82	0.79	0.83	0.84

Notes: This table presents the results of the VARX Model 2 shown in (5a), (5b), (5c) estimated over two periods: January 2015–November 2017 (35 observations) and December 2017–December 2019 (25 observations) based on either on the ask yields or the bid yields.

The exogenous variables are defined as in Table 5.

Standard errors of the coefficients are in parentheses and are based on the Newey and West's (1987) covariance matrix estimator with 6 lags.

* (resp. **) means that the coefficient is significant at 5% (resp. 1%) level.

Appendix A.

Table A1. Estimated coefficients for liquidity regressions (see Eq. (1)) using the full sample of green bonds and close yields

	Liquidity = BAS		Liquidity = ZTD		Liquidity = Roll	
	Coeff. (SE)	t-stat	Coeff. (SE)	t-stat	Coeff. (SE)	t-stat
<i>Constant</i>	0.0011 (0.0006)	1.63	0.0015 (0.0008)	1.82	0.0013 (0.0008)	1.62
<i>Liquidity</i>	0.082** (0.0198)	4.15	0.0006* (0.0003)	2.14	0.04027** (0.0134)	3.18
<i>Bondage</i>	-0.0002** (0.00004)	-5.83	-0.0002** (0.00006)	-3.55	-0.0002** (0.00006)	-3.52
<i>Amount issued</i>	-0.00165** (0.00015)	-10.93	-0.00172** (0.00016)	-10.85	-0.00172** (0.00016)	-10.81
<i>Coupon</i>	0.956** (0.004)	230.72	0.957** (0.006)	148.07	0.957** (0.006)	148.05
<i>TTM</i>	0.00002* (0.00001)	1.98	0.00003 (0.00001)	1.86	0.00003 (0.00001)	1.91
<i>Ratingscale</i>	0.00002 (0.00002)	1.24	0.00007 (0.00005)	1.49	0.00007 (0.00005)	1.52
<i>Term slope</i>	0.00031 (0.0002)	1.53	0.00025 (0.0002)	1.24	0.00026 (0.0002)	1.29
<i>Credit spread</i>	0.000001 (0.000003)	0.29	0.000001 (0.000005)	-0.25	0.000002 (0.00004)	0.52
<i>Certified</i>	0.0004** (0.00008)	4.58	0.00035* (0.00016)	2.08	0.00035* (0.00016)	2.20
<i>Currency</i>	0.00006 (0.00006)	1.06	-0.00006 (0.0002)	-0.29	-0.00007 (0.0002)	-0.32
<i>R²</i>	0.92		0.92		0.92	
<i>Adj. R²</i>	0.92		0.92		0.92	
<i>F-stat</i>	3347.1		3336.5		3339.9	
<i>Prob. (F-stat)</i>	0.000		0.000		0.000	

Note: For our full sample of green bonds, we run the liquidity regression set in Eq. (1) with the three different liquidity proxies (BAS, ZTD, Roll measure).

This table shows the coefficient estimates and their t-statistics (t-stats) for all variables presented in Eq. (1). Standard errors, which are corrected for time series effects, firm fixed effects, and heteroscedasticity are in parenthesis. Significance at 5% (*resp.* 1%) level of coefficients is marked * (*resp.* **).

Appendix B. Robustness checks

Table B1. Size of the liquidity premium (using Roll measure).

Panel A: Liquidity premium estimated according to the method of Dick-Nielsen (2012)

	Average (in bps)	Liquidity premium (in bps)			Number of observations		
		2-5y	5-10y	10-20y	2-5y	5-10y	10-20y
Full sample	11.7	12.6 (11.1 ; 14.1)	9.4 (8.3 ; 10.5)	14.2 (13.0 ; 15.4)	81	48	55
AAA	11.4	12.7 (11.2 ; 14.3)	9.5 (8.5 ; 10.5)	13.8 (12.6 ; 15.0)	60	28	33
AA and -	12.2	12.5 (10.9 ; 14.1)	9.2 (8.0 ; 10.4)	14.8 (13.5 ; 16.1)	21	20	22

Panel B: Liquidity premium estimated using the difference between CGB and SSAGB samples

	Average (in bps)	Liquidity premium (in bps)			Number of observations		
		2-5y	5-10y	10-20y	2-5y	5-10y	10-20y
Full sample	13.2	13.8 (12.5 ; 15.0)	12.3 (11.3 ; 13.3)	14.1 (13.0 ; 15.2)	81	48	55
AAA	12.9	13.5 (12.3 ; 14.7)	12.0 (11.1 ; 12.9)	13.9 (12.8 ; 15.0)	60	28	33
AA and -	13.8	14.4 (13.0 ; 15.8)	12.9 (11.7 ; 14.1)	14.2 (12.7 ; 15.7)	21	20	22

Notes: For every rating R we run the pooled regression: $\text{Yield}_{it} = \alpha R + \beta \text{Roll}_{it} + \text{controls}_{it} + \varepsilon_{it}$ where i refers to bond, t to time, and BAS_{it} is the bid-ask spread measure (denoted BAS).

Panel A displays the values of the liquidity premiums estimated according to the method of Dick-Nielsen (2012). We sort increasingly all values of BAS_{it} and find the median value BAS_{50} and the 5% value BAS_5 within each rating and maturity bucket (2-5y, 5-10y, and 10-20y). The liquidity component is defined as $\beta (\text{BAS}_{50} - \text{BAS}_5)$.

Panel B reports the values of liquidity premiums estimated using the difference between the median bid-ask spread of CGB sample and that of the SSAGB sample within each rating and maturity bucket (2-5y, 5-10y, and 10-20y). The liquidity component in the bucket is defined as $\beta (\text{BAS}_{50(\text{CGB})} - \text{BAS}_{50(\text{SSAGB})})$.

Both Panels A and B show for all buckets the liquidity premium component with standard errors in parenthesis. Confidence bands are found by a wild cluster bootstrap.