

# Tracing the Impact of the ECB's Asset Purchase Program on the Yield Curve\*

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We trace the impact of the European Central Bank's (ECB) asset purchase program (APP) on the yield curve. Exploiting granular information on sectoral asset holdings and ECB asset purchases, we construct a novel measure of the "free-float of duration risk" borne by arbitrageurs. We include this supply variable in an arbitrage-free term structure model in which central bank purchases reduce the free-float of duration risk and hence compress term premia. We estimate the stock of current and expected future APP holdings to reduce the 10-year yield by almost 1 percentage point. This reduction is persistent, with a half-life of five years.

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## 1. Introduction

We trace the impact of the European Central Bank's (ECB) asset purchase program (APP) on the euro area sovereign yield curve at announcement and over time. The ECB launched the APP in January 2015 by pledging the purchase of €60 billion of public and private sector securities a month from March 2015 until at least September 2016, amounting to €1.1 trillion. Successive rounds of recalibrations of the APP in December 2015, March 2016, December 2016, October 2017, and June 2018 took the eventual size of the portfolio to around €2.6 trillion by the end of net purchases in December 2018, corresponding to around 25 percent of euro area GDP. The ECB thus joined other major central banks, such as the Federal Reserve, in employing large-scale purchases—also known as “quantitative easing (QE)”—to provide monetary policy accommodation in the proximity of the effective lower bound by seeking to lower longer-term yields.<sup>1</sup>

For our analysis we deploy an affine term structure model in which central bank asset holdings compress term premia by reducing the amount of duration risk borne by arbitrageurs, building on Li and Wei (2013). In affine term structure models that are commonly used to study bond yield dynamics,<sup>2</sup> supply effects of securities do not play an explicit role. By contrast, the microfounded model by Vayanos and Vila (2021), featuring preferred-habitat investors and arbitrageurs, links the term premium to the amount of duration risk to be absorbed by the arbitrageurs: lower aggregate duration risk increases the risk-bearing capacity of the arbitrageurs, thereby decreasing risk compensation per unit of risk exposure (i.e., the “price of risk”) and hence the term premium. It is the overall amount of duration risk that matters for the term premium. Therefore, a change in bond supply at a specific maturity affects not only that

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<sup>1</sup>While the ECB previously embarked on outright asset purchases in the form of the Securities Markets Program (SMP), these purchases were categorically different from QE-type large-scale asset purchases and instead consisted of sterilized temporary interventions to provide liquidity to selected debt markets; see Eser and Schwaab (2016), Ghysels et al. (2016), and De Pooter, Martin, and Pruitt (2018).

<sup>2</sup>Cf. Kim and Wright (2005), Joslin, Singleton, and Zhu (2011), and Adrian, Crump, and Moench (2013).

maturity bracket but also term premia along the entire curve. Moreover, the model by Vayanos and Vila (2021) predicts that it is the whole sequence of current and discounted future aggregate duration in the market that determines current bond prices.

This link between bond supply and the term premium is captured in our term structure model by including a quantitative measure of duration risk in addition to standard level and slope yield curve factors. This allows us to study the term premium effect of the ECB's APP, which decreases the overall duration risk to be borne by arbitrageurs. Finally, as in Li and Wei (2013), we restrict the supply factor to not affect current and expected future short-term interest rates, thereby excluding a "rate signaling" channel of central bank asset purchases.<sup>3</sup>

Our empirical measure of duration risk in the market is inspired by the theory developed by Vayanos and Vila (2021). Rather than considering the exposure of all private investors, as in Li and Wei (2013) and Ihrig et al. (2018), we exploit security-level information on sectoral bond holdings from the ECB's Securities Holding Statistics (SHS) to develop a more granular measure. From total bond holdings we exclude not only bond holdings by domestic central banks and governments but also the portfolios of domestic hold-to-maturity investors as well as the foreign official sector, since these investor groups are unlikely to respond to changes in the supply and maturity structure of outstanding bonds. As a residual, we obtain the bond holdings of arbitrageurs. We weight these holdings

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<sup>3</sup>Signaling effects of the ECB's non-standard monetary policy measures have been analyzed by, e.g., Andrade et al. (2016), Arrata and Nguyen (2017), Lemke and Werner (2020), and Altavilla, Carboni, and Motto (2021). Such effects have been found to be small in magnitude compared with the effects of the duration extraction channel. By contrast, based on a shadow-rate term structure model estimated for the overnight index swap (OIS) yield curve, Geiger and Schupp (2018) find that unconventional monetary policy shocks have a stronger impact on expected short rates than on the forward term premium up to seven years. We do not separately identify the role of reserves creation for term premium compression as Christensen and Krogstrup (2019) do based on Swiss data, as reserve-and supply-induced effects cannot be independently identified for QE programs involving purchases of long-term securities. We also abstract from flow effects of purchases, which are of a more temporary nature; see D'Amico and King (2013) and Kandrac and Schlusche (2013) for the United States, Joyce and Tong (2012) for the United Kingdom, as well as Schlepper et al. (2017) and De Santis and Holm-Hadulla (2020) for the euro area.

according to their duration and normalize them by the total duration supply of outstanding government bonds. We refer to the share of duration risk exposure borne by arbitrageurs relative to total duration risk supply as the “free-float of duration risk.”

We estimate the model by minimizing the weighted sum of two fitting criteria. The first criterion measures the time-series fit of euro area sovereign bond yields (zero coupon, averaged across the largest four countries) over the period before markets started pricing large-scale asset purchases by the ECB. The second criterion is based on the fit of the cumulative yield decline over events (ECB press conferences and speeches) in the run-up to and around the announcement of the APP, which were perceived by markets to contain information on the forthcoming purchase program. We rely on this novel approach as our sample is relatively short: we can only construct our free-float measure from December 2009 based on the SHS data. Moreover, Eurosystem<sup>4</sup> bond holdings only became a significant source of variation in the free-float with the start of APP. This contrasts the U.S. experience, where the Federal Reserve’s monetary policy portfolio exhibited significant variation already before the inception of its large-scale asset purchases (LSAPs).

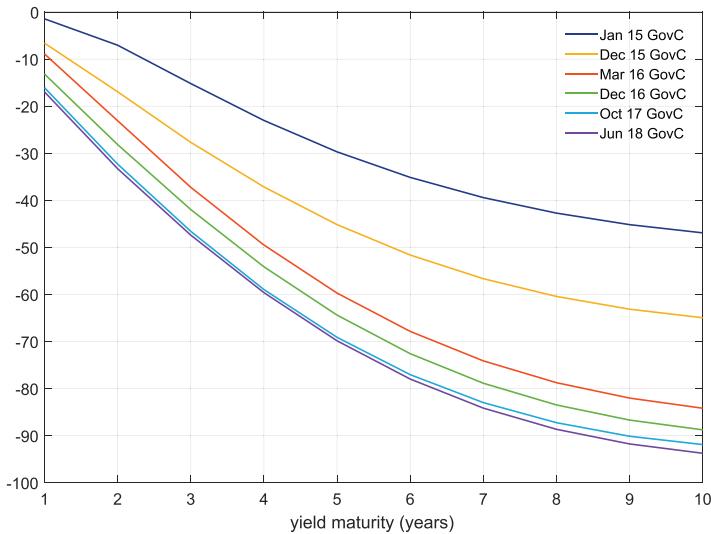
We report four main results. First, we find that the APP has flattened the yield curve and compressed term sovereign premia considerably; see Figure 1. Specifically, at its onset in January 2015, the APP compressed 10-year sovereign term premia by around 50 basis points (bps), the impact has increased gradually as the APP has been expanded in length and volume, and the impact has reached around 95 bps in June 2018. The 5–95 percent confidence interval, which accounts for parameter uncertainty, ranges from 65 to 130 bps.

Second, we find that the term premia compression due to the APP is persistent. Based on the path of APP net purchases envisaged by the Governing Council in June 2018, and assuming a horizon for full reinvestments of 3 years, we estimate a half-life of around 5 years for the 10-year term premium impact. The fading of the impact over time reflects, to some extent, the aging of the portfolio, i.e., the gradual loss of duration as the securities held in the portfolio mature,

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<sup>4</sup>The Eurosystem comprises the ECB and the 19 national central banks of the euro area member states.

**Figure 1. Impact of Different APP Vintages on the Sovereign Yield Curve**



**Note:** The figure shows the contemporaneous impact of the APP on the term structure of interest rates via the duration extraction channel. For the indicated dates  $t$  and maturities  $n$ , the respective point on the line provides an estimate of how much the sovereign  $n$ -period yield at the respective time  $t$  is compressed due to the impact on the term premium via the duration extraction channel.

as well as, in particular, the run-down of the portfolio that market participants anticipate to follow the reinvestment phase.

Third, the expected length of the reinvestment period after net purchases has a significant impact on term premia. The longer the reinvestment horizon, the larger the term premium impact. For example, under the counterfactual of no reinvestment, relative to an assumed reinvestment horizon of three years, the long-term interest rate would have been around 15 bps higher in June 2018.

Fourth, we use our model to make real-time predictions of the yield curve effect of the various APP recalibrations and compare these predictions with the observed yield curve reactions upon announcement, controlling for the expectations of APP parameters prevailing ahead of the announcement. We find that the model accounts well out of sample for the observed yield curve changes around APP announcements that implied major surprises regarding the future free-float.

Overall, our quantitative results for the yield impact of the APP (almost 50 bps in early 2015 and around 95 bps by mid-2018 for the 10-year big-four sovereign yield) are within the wide range of estimates reported in the literature. In comparing results, it is important to account for the different empirical approaches, as well as data and sample choices.<sup>5</sup> Several papers deploy event-study approaches, which focus (by design) on the surprise element of QE. While our econometric model also uses event information for the estimation, our impact estimates also exploit information on the full expected trajectory of ECB bond holdings. Quantifying the APP impact on euro area GDP-weighted average sovereign yields based on an event study, Altavilla, Carboni, and Motto (2021) find an impact of the initial APP package on the 10-year yield of 65 bps, while De Santis (2020) reports an APP impact on the 10-year yield of 72 bps over the period August 2014–October 2015 using a panel error correction specification with key APP dates informed by Bloomberg news stories. Bulligan and Delle Monache (2018) also conduct event-study regressions and find that news about the APP during September 2014 to July 2017 reduced the 10-year yield by around 50 bps. Andrade et al. (2016) conduct an event study over yields of individual sovereign bonds eligible under the APP and quantify the impact of the duration extraction channel of the initial APP package in early 2015 on bonds with a duration larger or equal to 10 years at 47 bps. Combining event information with survey-based evidence of investors' purchase expectations, Rostagno et al. (2021) estimate that APP has compressed the 10-year euro area sovereign yield by around 120 bps at the end of 2018. Blattner and Joyce (2020) consider an alternative GDP-scaled free-float measure and estimate, using a Bayesian vector autoregression (VAR), the impact of the

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<sup>5</sup>Koijen et al. (2021) estimate a demand system for government bonds by instrumental variables, finding that portfolio rebalancing actions of euro area investors between March 2015 and December 2017 reduced the average yield of the sovereign debt outstanding of each of the four largest euro area countries between 40 and 60 bps. As they average across maturities, their figures correspond roughly to our seven-year yield impact of 85 bps. Our findings are also consistent with the overview paper of Hartmann and Smets (2018), who report that in December 2016 the cumulated and joint impact of the APP together with credit-easing measures and interest rate cuts amounted to around 150 bps for the euro area 10-year sovereign yield.

initial APP package on the euro area 10-year bond yield at around 30 bps.

Our estimates are also broadly in line with those obtained for the U.S. Federal Reserve purchase programs, despite differences in market environment and purchase modalities on the two sides of the Atlantic. Applying the model by Li and Wei (2013) to the United States, Ihrig et al. (2018) estimate a peak cumulative impact of the Federal Reserve's LSAPs and its Maturity Extension Program of around 125 bps for a purchase volume of \$4.5 trillion. A direct comparison of our euro area peak impact estimates—around 95 bps in June 2018, with uncertainty bands ranging from 65 to 130 bps—with the U.S. figures is challenging due to factors such as a different sovereign bond market structure, a different global financial environment at the time of purchases, and a different allocation of purchases over time. Moreover, ideally a comparison would be based on a granular free-float measure of the type we construct, but its U.S. counterpart is not available to us.<sup>6</sup> Taking the size of the economy as a very rough yardstick of comparison, overall purchase volumes amount to around one-quarter of GDP in both cases. Hence, under such scaling, the U.S. impact would range in the upper part of our confidence band obtained for the APP.

The remainder of the paper is structured as follows. Section 2 explains the construction of our free-float measure and the yield data. Section 3 describes the model and inspects the mechanism of how central bank purchases affect the term premium. Section 4 outlines the estimation approach and documents the model fit. Section 5 reports the main results, i.e., the impact on the yield curve at different points in time, the persistence of those effects, the role of reinvestment, and the impact of selected APP recalibrations. It also sheds some light on the robustness of results. The last section concludes.

## 2. APP Duration Extraction and Euro Area Yields

We construct a theory-consistent measure of the free-float of duration risk, which enters our term structure model as supply variable

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<sup>6</sup>Using GDP as a scaling variable, as in Li and Wei (2013) and Ihrig et al. (2018), instead of total bond supply (as in our baseline specification) leaves our estimates largely unchanged, as discussed in more detail in Section 5.4.

(Section 2.1); we explain how to project it into the future using official ECB communication and survey information (Section 2.2); and we introduce the yield curve data (Section 2.3). Our analysis focuses on the government debt and average yields of the four largest euro area countries (Germany, France, Italy, and Spain; henceforth “big four”). These countries accounted for around 80 percent of euro area sovereign debt and around 76 percent of euro area GDP at the end of 2016.

## *2.1 A Theory-Consistent Measure of the Free-Float of Duration Risk*

As in Vayanos and Vila (2021), the term premium is affected by the amount of duration risk to be absorbed by arbitrageurs.<sup>7</sup> Motivated by their theoretical framework, we construct an empirical measure of the free-float of duration risk as follows:

$$\text{free-float of duration risk} = \frac{\text{duration-weighted bond holdings of arbitrageurs}}{\text{duration-weighted total bond supply}}. \quad (1)$$

The three key dimensions of this empirical measure of the free-float of duration risk are, first, which type of investors to count as arbitrageurs; second, the normalization of the free-float of duration risk with the total bond supply; and third, the range of securities considered for the measurement of duration risk.

Regarding the first dimension, we compute the duration-weighted bond holdings of arbitrageurs (the numerator in Equation (1)) in order to account for the role of different types of investors in the transmission of central bank asset purchases.<sup>8</sup> We divide

<sup>7</sup>See also Hamilton and Wu (2012). The appendix provides a more detailed discussion of the mapping between Vayanos and Vila (2021) and our model.

<sup>8</sup>To account for the duration of the bonds held by arbitrageurs, we consider the sectoral holdings in terms of their 10-year equivalents. The 10-year equivalent portfolio is a hypothetical portfolio that consists only of 10-year zero-coupon bonds and that has the same duration risk as the actual portfolio. The nominal amount (par value) of an individual bond  $j$  is converted into the 10-year equivalent using the following formula:  $10y\ equivalent_j = nominal_j \cdot \frac{duration_j}{10}$ . We use the maturity as a duration proxy. This measure has the advantage of

holding sectors into two groups—arbitrageurs and preferred-habitat investors—in line with Vayanos and Vila (2021). To distinguish the bond holdings of these two groups of investors, we exploit the granular information available in the Eurosystem Securities Holdings Statistics (SHS) on the sectors holding general government debt securities. The SHS data are available from 2009:Q4 at quarterly frequency. At the security level, these data provide information on the nominal value, the residual maturity, and the holding sectors. For euro area holdings, the SHS distinguishes the following holding sectors: monetary and financial institutions (MFI), money market funds (MMF), non-MMF investment funds, insurance corporations and pension funds (ICPF), other financial institutions, non-financial corporations (NFC), and households. By contrast, for foreign, i.e., non-euro area, holdings only a distinction between official and non-official portfolios is available.<sup>9</sup> In addition, we use information on

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abstracting from endogenous feedback effects from yield levels on other portfolio duration measures, such as modified duration. For a portfolio with weighted average maturity  $WAM = \sum_{j=1}^{bonds} (nominal_j \cdot maturity_j) / \sum_{j=1}^{bonds} nominal_j$ , the 10-year equivalent of the portfolio is obtained by weighting the portfolio's nominal value by  $\frac{WAM}{10}$ .

<sup>9</sup>The information on foreign holdings in the SHS is subject to two reporting biases, which we address as follows. First, nominal holdings by foreign private investors are inflated due to a custodial over-reporting bias of foreign non-official holdings. The SHS holdings for investors outside the euro area are collected from custodians. Custodians are financial institutions that hold securities on behalf of their customers. However, the custodians may not always know the final investor, especially if the customers of the custodians are institutions transacting on behalf of a third-party customer. If these third-party customers are located in the euro area, then the holdings reported by custodians as foreign holdings are in fact domestic holdings. As a result, the sum of all domestic and foreign holdings of some securities as reported in the SHS can exceed the outstanding amount of these securities as reported by the ECB's Government Finance Statistics (GFS). To address the over-reporting bias coming from custodians, we benchmark the nominal value of total outstanding government bonds for each country as reported in the SHS against the corresponding information from the GFS. We then adjust the foreign sector holdings obtained from the SHS downwards so that the sum of outstanding amounts across all sectors from the SHS data matches the totals from the GFS. Second, during the preliminary SHS data collection period from 2009:Q4 to 2013:Q3 foreign official sector holdings were largely unreported, as the submission of these data was voluntary. To address this issue, we backcast the nominal foreign official sector holdings for all countries from their 2013:Q4 levels based on the dynamics of euro area external financial liabilities as reported in the International Monetary Fund's Coordinated Portfolio Investment Survey (CPIS)

Eurosystem holdings derived from the ECB-internal security-level data on sovereign bond purchases.

Our group of preferred-habitat investors comprises the official sector, both euro area and foreign, and on the private sector side ICPFs. The inclusion of official holdings in the preferred-habitat category reflects the fact that these tend to have narrowly described mandates, which limit the degree to which these type of investors can engage in arbitrage. The official holdings comprise foreign exchange reserves by non-euro area central banks, holdings of the intra-euro area general government sector, as well as Eurosystem portfolios. The latter include both monetary-policy-related sovereign bond holdings, such as those accumulated under the Securities Markets Program (SMP) and the Public Sector Purchase Program (PSPP), as well as holdings which are unrelated to monetary policy and subject to the Agreement on Net Financial Assets (ANFA). Similarly, out of the private sector investors we include ICPFs in the preferred-habitat group, as these tend to follow hold-to-maturity strategies, matching long-dated liabilities with long-dated assets, and they are subject to regulatory requirements. ICPFs are, thus, unlikely or limited in their ability to rebalance away from their preferred habitats.

Our group of arbitrageurs is made up of all the private sectors other than ICPFs. These include—for the euro area—MFIs (excluding the Eurosystem), MMFs, non-MMF investment funds, NFCs, and households. In terms of magnitude, in the group of arbitrageurs MFIs are by far the dominant private domestic-holding sector of euro area sovereign bonds. In addition, we include the foreign non-official sector in the group of arbitrageurs.

We present some summary statistics of the asset holdings of the different sectors in Table 1. Before the start of the APP, euro area MFIs held the largest portion of big-four sovereign bonds, followed by the official sector other than the Eurosystem, which mainly reflects foreign reserve holdings. On balance, close to 55 percent of all outstanding big-four government bonds were in the hands of arbitrageurs pre-APP. Differences in the pre-APP average maturity

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and the Currency Composition of Official Foreign Exchange Reserves (COFER). In addition, we assume that the weighted average maturity (WAM) of foreign official sector holdings is constant over the preliminary SHS data collection period at the level of the average WAM of the pre-APP official reporting period.

**Table 1. Sovereign Bond Holdings by Investor Type and Sector**

	<b>Holdings (€bn)</b>		<b>WAM (Years)</b>	
	<b>Pre-APP</b>	<b>2018:Q2</b>	<b>Pre-APP</b>	<b>2018:Q2</b>
Arbitrageurs	3,281	2,632	6.1	6.7
MFI	1,334	1,141	5.0	5.4
Other Domestic	1,126	976	6.5	7.8
Foreign Non-official	815	514	7.3	7.7
Preferred-Habitat Investors	2,716	3,860	6.6	7.4
ICPF	1,009	1,241	10.8	10.6
Other Official	1,305	1,096	4.1	4.3
Eurosystem	402	1,523	4.5	7.0

**Note:** The table reports the nominal value and WAM of sector-specific holdings of bonds issued by the general governments of France, Germany, Italy, and Spain. The pre-APP period refers to average sector holdings in the period 2013:Q4 to 2014:Q4. The sector “Other Official” includes domestic euro area governments and the foreign official sector; “Other Domestic” includes NFCs, households, and financial institutions other than banks.

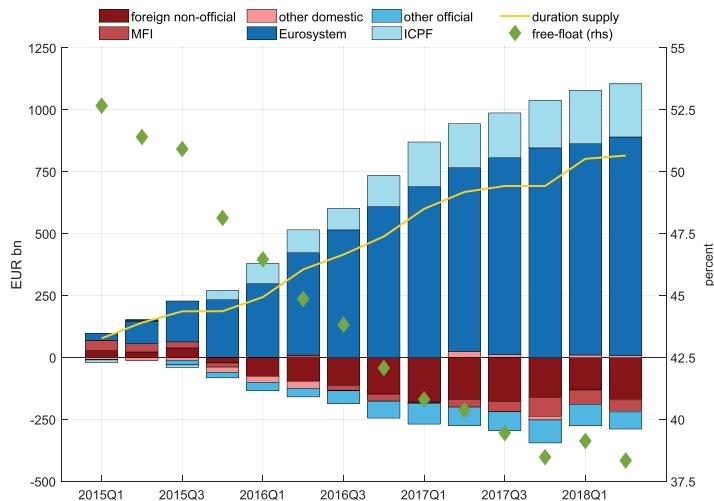
of sectoral portfolios point to different investment strategies. For instance, MFI holdings tended to be concentrated in shorter maturity segments compared with the maturity distribution of all outstanding government bonds, while ICPFs held substantially longer-dated paper.

Since the start of the APP, a notable portfolio rebalancing has taken place across sectors.<sup>10</sup> The share of Eurosystem holdings in total outstanding big-four government bonds has risen from less than 7 percent to around 23 percent by mid-2018; see Table 1. ICPFs were the only investors who increased their holdings alongside the Eurosystem. All sectors classified as arbitrageurs, and among these most prominently banks and foreign non-official investors, have been net sellers of government bonds. As a result, the share of holdings by arbitrageurs as a fraction of total outstanding big-four government bonds has fallen from 55 percent to around 41 percent.

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<sup>10</sup>See Bua and Dunne (2019) for a more detailed inspection deploying micro panel data from the investment fund industry domiciled and reporting in Ireland.

**Figure 2. Evolution of Duration Supply and Its Absorption by Investor Type**



**Note:** On the left axis, the chart shows cumulative changes in 10-year equivalent holdings of big-four general government bonds in €billion compared with pre-APP holdings as of 2014:Q4. On the right axis, the evolution of the free-float as a percentage of total duration supply is shown.

Figure 2 illustrates that since the start of the APP the Eurosystem has broadly offset the increase in the 10-year equivalent bond supply to the market through the issuance of big-four general government debt securities. ICPFs have increased their exposure to duration risk over the same period, which is consistent with their classification as preferred-habitat investors. By contrast, arbitrageurs, such as foreign private investors and, to a lesser extent, euro area banks, have reduced their relative exposure to duration risk since the start of the APP, as is evident from the material decline of the free-float.

Our classification of investor types as preferred habitat versus arbitrageurs is in line with the evidence on sectoral portfolio rebalancing in response to the PSPP in Koijen et al. (2017, 2021) and Bergant, Fidora, and Schmitz (2020). In particular, Koijen et al. (2017) and Koijen et al. (2021) identify the ICPF sector as the only one that increased its holdings euro area sovereign bonds—trading in the same direction as the ECB. The authors see the inelastic

or even upward-sloping demand by ICPFs reflecting their need to match their long-dated liabilities, which implies a preferred habitat for ICPF investments. By contrast, in particular the foreign sector and banks, but also mutual funds and households, decreased their holdings. Regarding the inclusion of MMFs, NFCs, and households in our group of arbitrageurs, we acknowledge that these groups may not arbitrage across assets in the same way that MFIs do. However, we see the fact that MMFs, NFCs, and households have rebalanced their portfolios in the same direction as MFIs and the foreign private sectors, as supporting the classification of MMFs, NFCs, and households as arbitrageurs. In any case, the holdings of MMFs, NFCs, and households are small in relative terms.

By way of comparison, most studies in the literature either take all government bonds in the hands of the private sector as the numerator of the relevant supply variable (in Equation (1)), or they account for foreign official bond holdings to some extent, while the preferred-habitat behavior of ICPFs is not accounted for. Specifically, in the U.S. context Li and Wei (2013) consider U.S. Treasury securities in the hands of all private investors as the numerator of the bond supply measure. Similarly, D'Amico et al. (2012) also work with the privately held Treasury supply. By contrast, foreign official holdings are accounted for by Hamilton and Wu (2012) who use, *inter alia*, the average maturity of outstanding debt, where the weights are given by the share of non-official-sector debt holdings in total debt of the respective maturity, whereas Kaminska and Zinna (2020) rely on a weighted average of the total Treasury bond supply, bond supply held by the Federal Reserve, and foreign bond holdings. In the European context, Blattner and Joyce (2020) deploy a measure of maturity-weighted debt, from which they subtract a proxy of bond holdings by the foreign official sector. Compared with the literature, we believe that by excluding both the official sector holdings and the holdings of ICPFs we obtain a better approximation of the holdings of "arbitrageurs" along the lines of Vayanos and Vila (2021).

Turning to the second dimension—the normalization of the duration-weighted bond holdings of arbitrageurs—we normalize by the total supply of duration risk, i.e., the 10-year equivalent value of the nominal amount of outstanding government bonds of the big-four euro area jurisdictions; see the denominator in Equation (1).

We refer to the share of duration risk exposure of arbitrageurs in the total duration risk supply as the free-float of duration risk; see the green diamonds in Figure 2. For the estimation of the model the quarterly free-float series is interpolated linearly to obtain observations at monthly frequency.

Whereas in the theoretical setting of Vayanos and Vila (2021) the quantities of bonds enter directly as market values, some scaling is necessary in an empirical context, for a growing bond market would in principle lead term premia (which are a function of bond supply) to increase without bound. Moreover, scaling can also be motivated by the fact that any bond-holding variable measured in nominal level terms is likely to be persistent and thus challenging for econometric inference; see also Kaminska and Zinna (2020).<sup>11</sup>

Normalizing by total bond supply captures investors' need to bear more risk when debt supply increases. At the same time, choosing the total bond supply may not fully account for investors' capacity to bear more risk in a growing economy. This may be captured to some extent by scaling by GDP, which is the approach taken by Li and Wei (2013), Greenwood and Vayanos (2014), and Blattner and Joyce (2020). In any case, in our robustness section (see Section 5.4), we investigate the sensitivity of our empirical estimates to the choice of normalizing variable and also present results obtained by using GDP as a scaling variable. We find that the results obtained using GDP as a scaling variable are very similar in terms of magnitude. This reflects the fact that the evolution of the total debt supply and GDP are correlated. Moreover, the similarity of the results obtained by using total bond supply and GDP as scaling values suggests that the impact estimates are mainly driven by the change of the free-float induced by central bank purchases, which dominate the slower-moving variations in the scaling variable.

With regard to the third dimension—the range of securities considered for our measure of the free-float of duration risk—we focus on the general government bonds of the four largest euro area countries (“big four”) purchased under the PSPP part of the APP. General government bonds comprise central government bonds, regional and local government bonds, as well as some social security funds.

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<sup>11</sup>Kaminska and Zinna (2020) use the amount of Treasuries held by arbitrageurs as scaling variable.

The APP initially consisted of three components: the PSPP, an asset-backed securities purchase program (ABSPP), and a covered bond purchase program (CBPP3). A fourth component was added with the corporate sector purchase program (CSPP) in March 2016. The PSPP is by far the largest component, making up 84 percent of total net purchases, against 8 percent in the CBPP3, 7 percent in the CSPP, and 1 percent in the ABSPP. Within the PSPP, 90 percent of purchases (88 percent until March 2016) are made in national sovereign bonds, while 10 percent (12 percent) are allocated to euro area supranational issuers. The allocation of purchases across national bond markets is guided by the subscription of the 19 euro area national central banks (NCBs) in the ECB's capital key.<sup>12</sup>

By focusing on the general government debt of the "big-four" euro area countries for our measure of duration risk, we abstract from the remaining 15 euro area countries which account for the remaining 20 percent of euro area debt, the purchase of other agencies and supranational bonds, as well as private sector purchases within the ABSPP, the CSPP, and the CBPP3. This aligns the range of securities included in our supply variable with the GDP-weighted yields of the "big-four" euro area countries, which is our variable of interest. While theory suggests that purchases of corporate bonds would likewise decrease the overall duration risk to be borne by the market and hence affect the term premium of government bonds, in practice (due to market fragmentation and the lower liquidity in the non-big-four bond markets) one may expect that the cross-market impacts of such purchases are more muted than those happening in the same issuer universe. Taking such differentiated effects into account would require incorporating a notion of market fragmentation into our model, which goes beyond the scope of this paper. Conversely, including non-big-four sovereign bonds and corporate bonds in the yield measure to align it with a supply measure that includes

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<sup>12</sup>The so-called ECB capital key refers to the subscription shares by the euro area NCBs in the capital of the ECB. The capital key subscription reflects the share of the respective member states in the total population and gross domestic product of the euro area, in equal measure.

duration risk from non-big-four sovereign bonds and corporate bonds would introduce non-negligible liquidity premia and credit risk premia into the yield measure. Moreover, importantly, our measure of the free-float of duration risk is a ratio, which captures the free-float of duration borne by arbitrageurs relative to total duration supply. Considering the private sector purchases and supply as well as the remaining 15 jurisdictions should leave the evolution of this ratio essentially unchanged, as both the numerator and the denominator of this ratio would adjust in that case. Finally, an advantage of our focus on the general government bonds of the big-four euro area countries is that for this set of securities it allows us to construct a granular and accurate measure of the free-float of duration risk. By contrast, using also the data for the remaining 15 countries, as well as private sector assets, would come at considerable computational cost and could introduce measurement errors, as the sectoral holdings data, which we would exploit to identify the holdings by arbitrageurs, require significant data cleaning.

## *2.2 Projecting APP Duration Extraction over Time*

Central bank asset purchases exert their impact on the term structure by reducing the free-float of duration risk to be borne by arbitrageurs. Importantly, the theoretical model by Vayanos and Vila (2021) implies that the yield impact of central bank asset purchases in a specific maturity spectrum depends on the evolution of the discounted duration of the stock of bonds held by the central bank over the entire life of bonds in this spectrum. Therefore, beyond measuring the contemporaneous free-float of duration risk, we also need to project, at any given point in time, the free-float of duration risk, and its reduction through central bank asset purchases, into the future.

Projecting the evolution of the duration-weighted central bank portfolio requires information on future purchase volumes. We use the fact that the ECB's forward guidance on the path of net asset purchases was communicated in terms of an intended monthly purchase pace and horizon. For example, at the initial announcement of the APP in January 2015, the ECB Governing Council communicated its intention to make net purchases of €60 billion a month from

March 2015 to at least September 2016. After this initial announcement, the Governing Council made changes to the purchase horizon and/or the size of monthly flows in December 2015, March 2016, December 2016, October 2017, and June 2018. Each of these dates provides an “APP vintage,” which is associated with a specific announced path for net purchases.<sup>13</sup>

In addition, we assume that announced net purchases are wound down along a linear tapering path, which reflects the ECB’s early guidance that net purchases would not end abruptly. The linear tapering is assumed to reduce the monthly net purchase volume from the announced end-date in steps of €10 billion. Moreover, in December 2015 it was announced that maturing principals would be reinvested “for as long as necessary.” From then on we assume a reinvestment phase to follow net asset purchases. From December 2015 to October 2017 we assume the reinvestment horizon to be two years. This is in line with median survey-based reinvestment expectation in the December 2017 Bloomberg survey, which first recorded reinvestment expectations. For June 2018, we use a median reinvestment horizon of three years as recorded in the respective Bloomberg survey. Table 2 summarizes, under the label “GovC,” the key parameters for the various APP vintages.

Projecting the duration-weighted central bank portfolio also requires information on the maturity distribution of purchases. As announced in January 2015, securities with maturity of 2 to 30 years were eligible for purchase. Within this spectrum, the Governing Council communicated that purchases would be made in a “market-neutral manner.” “Market neutrality” is understood to mean that the maturity distribution of the monthly flow of purchases is proportional to the eligible bond universe. Furthermore, initially, no purchases of securities with a yield below the deposit facility rate were undertaken. This constraint was relaxed in December 2016 from when purchases of securities with a yield below the deposit facility rate were allowed “to the extent necessary.” At the same point, the

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<sup>13</sup>The “time leg” of the forward guidance on asset purchases was complemented by a state-contingent forward-guidance element, according to which purchases would “in any case be conducted until the Governing Council would see a sustained adjustment in the path of inflation which is consistent with its aim of achieving inflation rates below, but close to, 2% over the medium term.”

**Table 2. APP Parameters Based on Governing Council Announcements and Survey Expectations**

Date	Type	Monthly Pace (€bn)	Horizon	Total Net Purchases (€bn)	WAM (Years)
Jan. 22, 2015	GovC	60 Linear Taper	Mar. 2015–Sep. 2016 Oct. 2016–Feb. 2017	1,290	8.29
Mar. 12, 2015	Survey	60 Linear Taper	Mar. 2015–Sep. 2016 Oct. 2016–Feb. 2017	1,290	8.43
Nov. 25, 2015	Survey	60 75 Linear Taper	Mar. 2015–Dec. 2015 Jan. 2016–Mar. 2017 Apr. 2017–Oct. 2017	1,988	8.30
Dec. 3, 2015	GovC	60 Linear Taper Reinvestment	Mar. 2015–Mar. 2017 Apr. 2017–Aug. 2017 Sept. 2017–Aug. 2019	1,650	8.23
Mar. 3, 2016	Survey	60 80 Linear Taper Reinvestment	Mar. 2015–Mar. 2016 Apr. 2016–Mar. 2017 Apr. 2017–Oct. 2017 Nov. 2017–Oct. 2019	2,020	8.69
Mar. 10, 2016	GovC	60 80 Linear Taper Reinvestment	Mar. 2015–Mar. 2016 Apr. 2016–Mar. 2017 Apr. 2017–Oct. 2017 Nov. 2017–Oct. 2019	2,020	8.70
Dec. 2, 2016	Survey	60 80 72 Linear Taper Reinvestment	Mar. 2015–Mar. 2016 Apr. 2016–Mar. 2017 Apr. 2017–Aug. 2017 Sep. 2017–Apr. 2018 May 2018–Apr. 2020	2,388	8.49

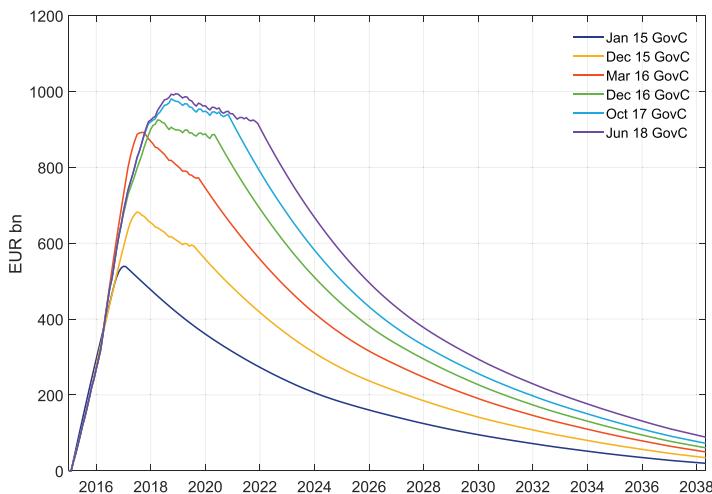
(continued)

**Table 2. (Continued)**

Date	Type	Monthly Pace (€bn)	Horizon	Total Net Purchases (€bn)	WAM (Years)
Dec. 8, 2016	GovC	60 80 60	Mar. 2015–Mar. 2016 Apr. 2016–Mar. 2017 Apr. 2017–Dec. 2017 Jan. 2018–May 2018 Jun. 2018–May 2020	2,430	7.27
Oct. 18, 2017	Survey	60 80 60	Mar. 2015–Mar. 2016 Apr. 2016–Mar. 2017 Apr. 2017–Dec. 2017 Jan. 2018–Sep. 2018 Oct. 2018–Sep. 2020	2,550	7.52
Oct. 26, 2017	GovC	60 80 60 30	Mar. 2015–Mar. 2016 Apr. 2016–Mar. 2017 Apr. 2017–Dec. 2017 Jan. 2018–Sep. 2018 Oct. 2018–Nov. 2018 Dec. 2018–Nov. 2020	2,580	7.21
Jun. 7, 2018	Survey	60 80 60 30	Mar. 2015–Mar. 2016 Apr. 2016–Mar. 2017 Apr. 2017–Dec. 2017 Jan. 2018–Sep. 2018 Oct. 2018–Dec. 2018 Jan. 2019–Dec. 2021	2,593	7.24
Jun. 14, 2018	GovC	60 80 60 30 15	Mar. 2015–Mar. 2016 Apr. 2016–Mar. 2017 Apr. 2017–Dec. 2017 Jan. 2018–Sep. 2018 Oct. 2018–Dec. 2018 Jan. 2019–Dec. 2021	2,595	7.24

**Note:** The table reports the main APP purchase parameters for the different Governing Council announcements ("GovC"), as well as based on market expectations as reflected in Bloomberg surveys ("Survey"). The WAM reported is the WAM at the end of the net purchases associated. For each GovC series, a linear tapering is added to the announced pace and horizon, in line with the GovC communication that net purchases would not end abruptly. The survey parameters reported reflect the median responses. For Mar. 2015, Nov. 2015, and Mar. 2016 only the expected end date of net purchases is reported in the survey, and the linear tapering is added by assumption. Following the announcement at the Dec. 2015 GovC that maturing principals would be reinvested "for as long as necessary," a reinvestment phase is assumed. From Dec. 2015 to Oct. 2017, the assumed reinvestment phase is two years, in line with median survey-based reinvestment expectation in the Dec. 2017 survey, which first recorded such expectations. In Jun. 2018, the median reinvestment expectation was three years.

**Figure 3. Evolution of ECB's Duration-Weighted Government Bond Holdings**



**Note:** The figure shows the projected evolution of the government bond holdings for the big-four euro area countries in terms of 10-year equivalents. The different paths correspond to the APP vintages summarized in Table 2.

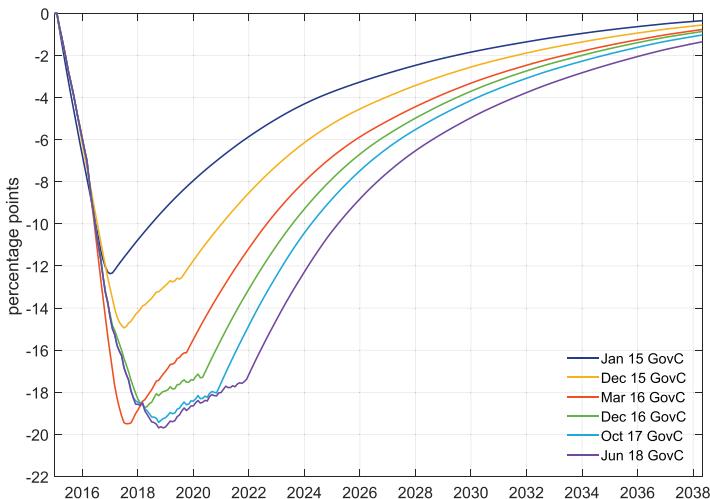
eligible maturity spectrum was extended from an interval spanning 2 to 30 years to a wider range of 1 to 30 years.

In addition, the projections account for further eligibility and operational criteria, which guided the implementation of historical purchases and affect their composition along several dimensions. First, the distribution of purchases across countries is determined by the ECB's capital key. Second, purchases respect an issue and issuer limit of 33 percent. Finally, future issuance of purchaseable securities is taken into account and based on the debt projections that enter the ECB's quarterly staff macroeconomic projection exercises and that were available at the given point in time.<sup>14</sup> Figure 3 shows the resulting projections of the APP portfolio in terms of 10-year equivalents for the different APP vintages summarized in Table 2.

Finally, we construct the trajectory of the free-float of duration risk over time by complementing the projections of the Eurosystem

<sup>14</sup>For details on the aggregate debt projections, see Bouabdallah et al. (2017).

**Figure 4. Reduction in the Free-Float  
for Different APP Vintages**



**Note:** The figure shows the compression of the free-float measure induced over time by the successive vintages of the APP. Each line shows the reduction of the free-float relative to the counterfactual of no APP. The different vintages are summarized in Table 2.

duration absorption with projections for duration supply. The projected duration supply is again based on the debt projections that enter the Eurosystem staff projections at a given point and are hence revised over time. In addition, we make the assumption that the WAM of the market portfolio remains unchanged over the projection horizon at the last observed WAM in each bond market. Figure 4 illustrates the compression of the free-float measure induced over time by the different vintages of the APP, constructed under the assumption that the ECB purchases reduce exclusively the holdings of price-sensitive investors. Each line shows the reduction of the free-float relative to the counterfactual of no APP.

Below we also study the effect of announcements on asset purchases on the yield curve. These should have an impact on sovereign bond yields only to the extent that they are unanticipated by financial markets. To quantify the yield impact of the initial APP announcement, as well as subsequent recalibration vintages, we therefore isolate the surprise in terms of additional future duration

absorption associated with each announcement relative to what is already priced in based on pre-announcement market expectations, similar to Ihrig et al. (2018). We exploit the regular surveys by Bloomberg to obtain market expectations on the future purchase path of the APP. The resulting parameters are summarized in Table 2 under the label “Survey.” We show for every APP recalibration the corresponding market expectations ahead of the recalibration announcement. In addition, we report the March 2015 survey path, as we use this in the estimation of our model (see Section 4.1).

The Bloomberg surveys were conducted systematically every six weeks from March 2015, and are typically published in the days ahead of the ECB Governing Council meetings.<sup>15</sup> The March 2015, November 2015, and December 2016 surveys did not contain information on expected “tapering” volumes. In those cases, we assume a linear tapering. The December 2016 survey contained information on expected tapering volumes, which we take into account. The October 2017 and June 2018 surveys provided a fully specified path for net asset purchases. Starting from the Governing Council’s December 2015 reinvestment announcement until October 2017, we use a two-year reinvestment phase for the survey-based APP projections, in line with the Bloomberg survey of December 2017. For June 2018, we use a three-year reinvestment horizon in line with the corresponding survey. The maturity distribution of purchases for the survey vintages is assumed to follow the market-neutrality principle, in line with the approach taken for the Governing Council vintages. Using the survey-based information on the expected path of the APP, as well as assuming a market neutral maturity distribution of purchases, we create projections of the evolution of the market-expected duration-weighted APP portfolio, and the implied reduction of the free-float of duration risk.

### 2.3 Yields

In contrast to the U.S. Treasury market, there is no single sovereign fixed-income market at the level of the euro area as a whole.

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<sup>15</sup>The March 2015 Governing Council meeting and survey provide an exception: the March 12, 2015 survey was conducted and published after the March 5, 2015 Governing Council meeting.

Each individual sovereign issues its own bonds. In order to provide a good representation of the overall sovereign debt market of euro area countries, we focus on the dynamics of the synthetic big-four euro area sovereign yield curve, which we construct as the GDP-weighted average of zero-coupon yields of Germany, France, Italy, and Spain.<sup>16</sup> The country-specific zero-coupon yields are constructed from prices of nominal bonds reported on the MTS platform. Bond prices are converted to zero-coupon yields using the Nelson-Siegel-Svensson methodology.<sup>17</sup> Our econometric analysis starts in December 2009 (in line with the availability of our SHS data) and ends in June 2018, when the Governing Council first expressed its anticipation to cease asset purchases by the end of December 2018, which was then subsequently confirmed. Figure 5 shows daily time series of the synthetic big-four zero-coupon yields for selected maturities over this period.

### 3. The Model

#### 3.1 A Term Structure Model with Quantities

For tracing the impact of the APP on the yield curve, we rely on the model introduced by Li and Wei (2013). Yield curve dynamics are parsimoniously captured by three observable factors. The first two factors are given by the first two principal components (PCs) extracted from a cross-section of observed yields; see details in Section 4.1. We denote the first PC as the level factor  $L_t$  and the second PC as the slope factor  $S_t$ . The third factor,  $Q_t$ , is our free-float measure; see Equation (1). We collect the three factors in the vector  $X_t = (L_t, S_t, Q_t)'$ .

The short-term interest rate  $i_t$  is a linear combination of the factors

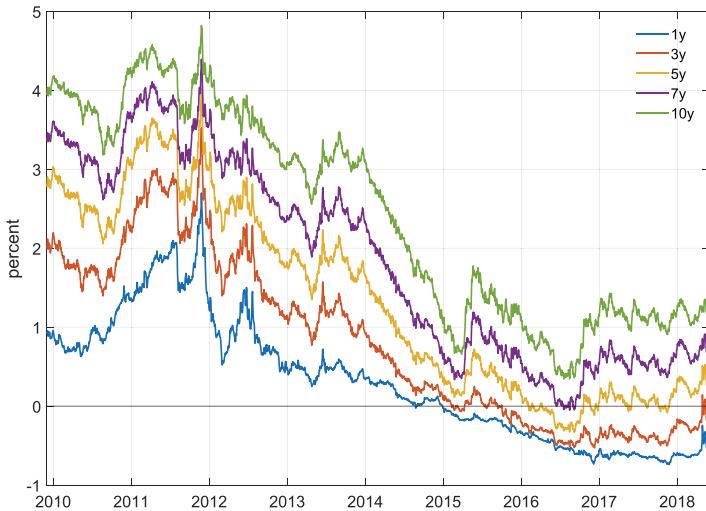
$$i_t = \delta_0 + \delta'_1 X_t, \quad (2)$$

where we impose the constraint that  $\delta_1 = (\delta_{1L}, \delta_{1S}, 0)'$ , i.e., as in Li and Wei (2013),  $Q_t$  does not affect the short rate contemporaneously.

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<sup>16</sup>The weights for Germany, France, Italy, Spain, are 0.38, 0.27, 0.21, 0.14, respectively.

<sup>17</sup>Information about the ECB's methodology for deriving zero-coupon yields is available at [https://www.ecb.europa.eu/stats/financial\\_markets\\_and\\_interest\\_rates/euro\\_area\\_yield\\_curves/html/index.en.html](https://www.ecb.europa.eu/stats/financial_markets_and_interest_rates/euro_area_yield_curves/html/index.en.html).

**Figure 5. Euro Area Zero-Coupon Yields**

**Note:** The figure displays daily time series of the GDP-weighted synthetic zero-coupon yields for selected maturities of the big-four euro area countries for selected maturities from December 2009 to June 2018.

The factors  $X_t$  follow a VAR(1),

$$X_t = c + \mathcal{K}X_{t-1} + \Omega\epsilon_t, \quad \epsilon_t \sim N(0, I). \quad (3)$$

Following Li and Wei (2013), we constrain the autoregressive matrix  $\mathcal{K}$  to be block-diagonal with the two blocks  $(L_t, S_t)$  and  $Q_t$ . In addition, the contemporaneous shock impact matrix  $\Omega$  is assumed to be lower triangular:

$$\mathcal{K} = \begin{pmatrix} \mathcal{K}_{LL} & \mathcal{K}_{LS} & 0 \\ \mathcal{K}_{SL} & \mathcal{K}_{SS} & 0 \\ 0 & 0 & \mathcal{K}_{QQ} \end{pmatrix}, \quad \Omega = \begin{pmatrix} \Omega_{LL} & 0 & 0 \\ \Omega_{SL} & \Omega_{SS} & 0 \\ \Omega_{QL} & \Omega_{QS} & \Omega_{QQ} \end{pmatrix}. \quad (4)$$

Together with the assumption that the last element of  $\delta_1$  is zero, Equation (4) implies that the free-float measure  $Q_t$  does not forecast the short rate. In other words, the model excludes a potential

signaling channel of central bank asset purchases.<sup>18</sup> In addition, the quantity measure is assumed not to be predictable by the yield curve factors.

The pricing kernel  $M_t$  is exponentially affine in the factors

$$M_{t+1} = \exp(-i_t - 0.5\lambda'_t\lambda_t - \lambda'_t\epsilon_{t+1}), \quad (5)$$

where the market prices of risk  $\lambda_t$  are also affine in the factors

$$\lambda_t = \lambda_0 + \Lambda_1 X_t. \quad (6)$$

As in Li and Wei (2013) we impose the following zero constraints on the risk-compensation parameters  $\lambda_0$  and  $\Lambda_1$ :

$$\lambda_0 = \begin{pmatrix} \lambda_{0,L} \\ \lambda_{0,S} \\ 0 \end{pmatrix}, \quad \Lambda_1 = \begin{pmatrix} \Lambda_{1,LL} & \Lambda_{1,LS} & \Lambda_{1,LQ} \\ \Lambda_{1,SL} & \Lambda_{1,SS} & \Lambda_{1,SQ} \\ 0 & 0 & 0 \end{pmatrix}. \quad (7)$$

This means that we assume that only level and slope risk is priced, but that the corresponding risk prices are driven by all three factors, including the quantity variable. Innovations to the quantity variable themselves are not priced, i.e., their market price of risk is zero.<sup>19</sup>

The market price of risk vector  $\lambda_t$  is where the effects of central bank asset purchases are determined in the model: changes in the quantity variable affect risk prices of level and slope risk and thereby term premia. The economic interpretation of this link between bond supply and term premia is discussed in more detail in Section 3.3 below.

Zero-coupon bond prices  $P_t^n$  of bonds with maturity  $n$  satisfy the no-arbitrage pricing equation

$$P_t^n = E(M_{t+1} P_{t+1}^{n-1} | X_t), \quad P_t^0 = 1. \quad (8)$$

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<sup>18</sup>We worked with specifications that relax the zero restrictions in the short-rate equation (2) and the autoregressive matrix in (3) in order to allow for an impact of the free-float on (physical) short-rate expectations and hence for a signaling channel of QE to be active. However, those specifications did not lead to meaningful results (they were either insignificant or wrong-signed), so we did not pursue this route further.

<sup>19</sup>Li and Wei (2013) argue that “Treasury supply is unlikely to be a source of undiversifiable risk that should be priced on its own.” Moreover, the zero restrictions on  $\lambda_0$  and  $\Lambda_1$  help to establish the analytical solutions for bond prices under given (“perfect foresight”) investor expectations of future bond holdings as introduced in the following section.

Bond prices are converted into yields via  $y_t^n = -\frac{1}{n} \log P_t^n$ . Given the affine structure of the model, yields turn out to be affine functions of  $X_t$ ,

$$y_t^n = -\frac{1}{n} (A_n + B'_n X_t), \quad (9)$$

where  $A_n$  and  $B_n$  satisfy the usual difference equations

$$A_{n+1} = A_n + B'_n (c - \Omega \lambda_0) + \frac{1}{2} B'_n \Omega \Omega' B_n - \delta_0, \quad (10)$$

$$B'_{n+1} = B'_n (\mathcal{K} - \Omega \Lambda_1) - \delta'_1, \quad (11)$$

with  $A_0 = 0$  and  $B_0 = 0$ . If  $\Lambda_{1,LQ} \neq 0$  and  $\Lambda_{1,SQ} \neq 0$ , then  $B_{n,Q}$ , the third element of  $B_n$ , is different from zero. Hence, the quantity variable affects bond yields through the impact on the risk-compensation parameters but does not have a contemporaneous effect on the short rate.<sup>20</sup>

### 3.2 Bond Supply and the Yield Curve: Modeling Anticipated Shocks

How does the model translate a change in central bank asset purchases into changes in term premia and bond yields?

Under the standard approach in affine term structure models, time- $t$  bond yields can only change if the time- $t$  factors change. Iterating the no-arbitrage bond pricing Equation (8) forward, one obtains

$$P_t^n = E(M_{t+1} \cdot M_{t+2} \cdot \dots \cdot M_{t+n} | X_t). \quad (12)$$

Bond prices (and thus yields) depend on the expected sequence of future pricing kernels (short rates and risk compensation), which

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<sup>20</sup>The quantity variable does forecast level and slope factors—and hence future short-term rates via (2)—under the so-called risk-neutral dynamics. In a nutshell, under the risk-neutral probability measure, denoted by  $\mathbb{Q}$ , the pricing equation (8) boils down to discounting with the short-term rate (as opposed to the pricing kernel (5) involving risk corrections), i.e.,  $P_t^n = E^{\mathbb{Q}}(e^{-it} P_{t+1}^{n-1} | X_t)$ , and incorporating the risk correction implicitly by using an amended factor process, which under the  $\mathbb{Q}$  measure follows  $X_t = \tilde{c} + \tilde{\mathcal{K}} X_{t-1} + \Omega \epsilon_t^Q$ , where  $\tilde{c} = c - \Omega \lambda_0$  and  $\tilde{\mathcal{K}} = \mathcal{K} - \Omega \Lambda_1$ .

are, in turn, a function of future state variables and their innovations. Thus, current yields depend on the full path of our free-float measure over the lifetime of the bond. At the same time, the information set in Equation (12) is just the current state variables  $X_t$  which are driven by a VAR. As a result, expectations of future state variables (including  $Q$ ) and pricing kernels can only change if the current state variables  $X_t$  change. Accordingly, as indicated by the closed-form solution of the model (9), bond yields change if and only if current factors change. In particular, changes to central bank asset purchases could only be captured by a change to current  $Q_t$ . This change would trigger a change of future expected  $Q_{t+h}$  and future risk pricing—via Equation (6)—but this change in expectations of future quantities is fully determined by the change in the current (time- $t$ ) quantity.

However, the rigid link between future expectations and current state variables—implied by the standard approach described before—does not square well with the empirical evidence. In practice, the pure announcement of central bank asset purchases (i.e., statements affecting future  $Q_{t+h}$ ) can have a significant impact on the yield curve today without contemporaneously moving  $Q_t$  at all. Moreover, while asset purchases are ongoing, too, further announced and credible changes to future purchase parameters—for example, a prolongation of the reinvestment horizon—can affect current bond yields even if they do not contemporaneously affect  $Q_t$ . Finally, even if an innovation to the path of central bank asset purchases does affect the current free-float  $Q_t$ , the announced future changes to  $Q$  may differ from those implied by the conditional expectations  $E_t(Q_{t+h})$  as prescribed by the VAR in Equation (3).

To capture the possibility that anticipated future free-float changes have an impact on the current yield curve over and beyond what is implied by the current states, Li and Wei (2013) allow anticipated innovations to the quantity variable to enter the bond pricing equation. Specifically, their approach amounts to conditioning bond pricing not only on current state variables, as in Equation (8), but also on a sequence of anticipated future free-float ratios  $\bar{Q}_t = \{\bar{Q}_t, \bar{Q}_{t+1}, \bar{Q}_{t+2}, \dots\}$

$$P_t^n = E(M_{t+1} P_{t+1}^{n-1} | X_t, \bar{Q}_t), \quad P_t^0 = 1. \quad (13)$$

This is the same pricing expression as in Equation (8), except that it uses an enhanced set of conditioning information. Denote by  $\mathcal{Q}_t^0 = \{Q_t^0, Q_{t+1}^0, Q_{t+2}^0, \dots\}$  the sequence of expected free-float ratios based on the state vector  $X_t$  and the VAR dynamics in Equation (3). Let  $u_{t+h} = \bar{Q}_{t+h} - Q_{t+h}^0$  denote the anticipated innovation of the free-float to the “baseline” and  $\mathcal{U}_t = \{u_t, u_{t+1}, u_{t+2}, \dots\}$  the corresponding sequence of anticipated innovations. Li and Wei (2013) show that bonds priced under the enhanced information set in Equation (13) satisfy the yield equation:

$$y_t^n = -\frac{1}{n} A_n + d y_n(\mathcal{U}_t) - \frac{1}{n} B'_n X_t, \quad (14)$$

where  $A_n$  and  $B_n$  are the same expressions as in Equation (9) in the standard setup and

$$d y_n(\mathcal{U}_t) = -\frac{1}{n} \left[ B_{n,Q} u_t + \sum_{h=1}^n B_{n-h,Q} (u_{t+h} - \mathcal{K}_{QQ} u_{t+h-1}) \right]. \quad (15)$$

The expression  $d y_n(\mathcal{U}_t)$  is the impact of a sequence of anticipated innovations to the quantity factor on the  $n$ -period term premium and corresponding yield over and beyond what is incorporated in current factors  $X_t$ .<sup>21</sup>

We can alternatively rearrange the terms in Equation (15) to obtain

$$d y_n(\mathcal{U}_t) = \sum_{h=1}^n \gamma_h^n u_{t+h-1} = \gamma^{n'} \mathcal{U}_t, \quad (16)$$

where  $\gamma^n = (\gamma_1^n, \dots, \gamma_n^n)'$  with

$$\gamma_h^n = -\frac{1}{n} (B_{n-h+1,Q} - \mathcal{K}_{QQ} B_{n-h,Q}). \quad (17)$$

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<sup>21</sup>We consistently use  $\mathcal{U}$  sequences that are longer than the lifetime of any bond. Therefore, we do not need to obey the distinction in Li and Wei (2013) regarding the upper summation limit that becomes relevant if  $\mathcal{U}$  sequences are shorter than bond maturities. In terms of notation, in any scalar product involving  $\mathcal{U}$ , such as, e.g., in (16),  $\mathcal{U}$  is assumed to have the same length as the corresponding multiplying vector.

The model stipulates a linear relationship between changes in the trajectory of the anticipated future free-float over the tenor of a bond and the change in the yield of that bond. The sensitivity of yields to anticipated innovations in the future free-float are captured by maturity- and horizon-specific “impact factors”  $\gamma_h^n$ , which are a function of the model parameters, such as the persistence of factors, innovation volatility, and market prices of risk.

In Section 5 we deploy Equation (16) to investigate how APP recalibrations have affected the yield curve. A certain APP surprise at time  $t$  is summarized by a corresponding  $\mathcal{U}_t$  sequence and the yield impact is obtained via (16).

### 3.3 Bond Supply and the Yield Curve: Interpreting the Transmission Channel

In this section we provide further detail on the transmission channel of shocks to the free-float of duration risk implied by central bank asset purchases in our empirical model. First, we recall the standard decomposition of yields and show how expected future free-float measures  $E(Q_{t+h}|X_t)$ —with expectations being fully determined by current states—affect expected future excess returns and hence term premia. Second, we show that the same transmission channel holds for anticipated shocks, i.e., free-float innovations  $u_{t+h}$  that are *not* implied by contemporaneous state variables: we find that the effect of such an anticipated free-float shock  $u_{t+h}$  on future expected excess returns and hence term premia is the same as the effect stemming from a change in expected free-float induced by a change in current states, i.e.,  $E(Q_{t+h}|X_t)$ .

The  $n$ -period bond yield can be represented as the sum of the expectations component (average expected future short rates over the lifetime of the bond) and the term premium. The term premium component is, in turn, given by the average of expected future excess returns:

$$y_t^n = \underbrace{\frac{1}{n} E_t \sum_{h=0}^{n-1} i_{t+h}}_{\text{Expectations component}} + \underbrace{\frac{1}{n} E_t \sum_{h=1}^n r x_{t+h}^{n-h}}_{\text{Term premium}}, \quad (18)$$

where  $rx_{t+h}^{n-h} = \ln P_{t+h}^{n-h} - \ln P_{t+h-1}^{n-h+1} - i_{t+h-1}$  is the one-period excess return for a bond with maturity  $n-h+1$  purchased at time  $t+h-1$ . Unless specified otherwise, the conditional expectation  $E_\tau(\cdot)$  is equivalent to  $E(\cdot|X_\tau)$ .

The identity in (18) is independent of a specific model. Different term structure models imply different parametric expressions for the expectations component and the term premium. For the affine model introduced in Section 3.1, each expected future excess return, conditional on information at the time of the purchase of the bond, can be expressed as<sup>22</sup>

$$E_{t+h-1} rx_{t+h}^{n-h} = B'_{n-h} \Omega \lambda_{t+h-1} + JI. \quad (19)$$

The term  $B'_{n-h} \Omega$  captures factor sensitivity or “duration risk,” i.e., the exposure of (log) bond prices to unexpected changes in risk factors, while  $\lambda_{t+h-1}$  is the time-varying “price of risk,” i.e., the amount of excess return compensation per unit of risk. This compensation varies over time but is the same across bonds of all maturities, thus excluding arbitrage opportunities. The last item is a convexity adjustment (Jensen inequality) term, given by  $JI = -0.5 B'_{n-h} \Omega \Omega' B_{n-h}$ , which does not depend on the factors.

The zero restrictions in (7) imply that level and slope risk is priced, while the risk at any future time of unexpected changes in the free-float measure is not priced, i.e.,  $\lambda_{t+h-1,Q} \equiv 0$ . Grouping the level and slope factor as  $Z_t = (L_t, S_t)'$  and writing the corresponding model matrices in a partitioned fashion (denoting the upper  $2 \times 2$  part of  $\Lambda_1$  in Equation (7) by  $\Lambda_{1,ZZ}$ , etc.), the market price of level/slope risk, i.e., the  $2 \times 1$  vector  $\lambda_{t+h-1,Z}$ , is given by

$$\lambda_{t+h-1,Z} = \lambda_{0,Z} + \Lambda_{1,ZZ} Z_{t+h-1} + \Lambda_{1,ZQ} Q_{t+h-1}. \quad (20)$$

The time variation in the market price of level/slope risk is driven by the level and slope itself ( $\Lambda_{1,ZZ} Z_{t+h-1}$ ) as well as by the free-float measure ( $\Lambda_{1,ZQ} Q_{t+h-1}$ ). Rewriting (19) we obtain

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<sup>22</sup>The link between Equations (18) and (19) can be seen by conditioning, in (19), the future expected one-period excess return on information (factors) at time  $t$  and applying the law of iterated expectations.

$$\begin{aligned}
E_{t+h-1}rx_{t+h}^{n-h} &= \underbrace{\Theta_{n-h} + \theta'_{n-h}Z_{t+h-1}}_{\text{Terms independent of } Q} \\
&+ \underbrace{(B'_{n-h,Z}\Omega_{ZZ} + B_{n-h,Q}\Omega_{QZ})}_{\text{Exposure to level/slope risk}} \underbrace{\Lambda_{1,ZQ}Q_{t+h-1}}_{\text{Price of risk components driven by } Q}, \tag{21}
\end{aligned}$$

where  $\Theta_{n-h}$  is a constant comprising the Jensen term  $JI$  in (19) and the time-invariant risk compensation (as function of  $\lambda_{0,Z}$ ). The term  $\theta'_{n-h}Z_{t+h-1}$  depends on the level and slope factors but not on the free-float measure. The last summand is the time-varying contribution of the quantity factor to the one-period expected excess return  $h-1$  periods ahead. The first part in parentheses (“Exposure to level/slope risk”) is the log bond price sensitivity to  $\epsilon_Z = (\epsilon_L, \epsilon_S)'$  shocks in (3). This part affects the level and slope factors, and hence bond prices, either directly, via  $B'_{n-h,Z}\Omega_{ZZ}$ , or indirectly, by contemporaneously affecting the  $Q$  factor (via  $\Omega_{QZ}$ ) and affecting bond prices via the respective factor loading  $B_{n-h,Q}$ . The second part (“Price of risk components driven by  $Q$ ”) is the time-varying contribution of  $Q_{t+h-1}$  to the respective prices of level and slope risk. Conditioning Equation (19) on information at time  $t$ , we note that if the current free-float  $Q_t$  changes, this affects  $E_t(Q_{t+h-1})$  via the VAR, which in turn shifts expected future excess returns at time  $t$  through a change in the expected market price of risk and thus the time- $t$  term premium.

Having shown how expected free-floats—with expectations spanned by current state variables—have an impact on term premia in the standard approach of affine term structure models, we now explain that the same transmission channel holds for anticipated shocks, i.e., free-float innovations  $u_{t+h}$  that are not implied by contemporaneous state variables.

Recall that the impact factors  $\gamma_h^n$  in Equation (16) are expressed in terms of factor loadings on the free-float factor  $B_{\cdot,Q}$ ; see Equation (17). We now convert them into an alternative expression that highlights their economic interpretation as risk premium contribution. Starting from the recursion in Equation (11) and defining a selection vector  $s = (0, 0, 1)'$ , the impact of  $Q$  on the  $m$ -maturity log bond price is given by

$$B_{m,Q} = B'_{m-1}\mathcal{K}s - B'_{m-1}\Omega\Lambda_1 s - \delta'_1 s. \tag{22}$$

Grouping again  $Z_t = (L_t, S_t)'$ , partitioning system matrices accordingly, and noting the zero restrictions in  $\mathcal{K}$ ,  $\Omega$ , and  $\Lambda_1$ , we obtain

$$\mathcal{K} \cdot s = \mathcal{K}_{QQ} \cdot s, \quad \Omega \Lambda_1 s = \begin{pmatrix} \Omega_{ZZ} \Lambda_{1,ZQ} \\ \Omega_{QZ} \Lambda_{1,ZQ} \end{pmatrix}, \quad \delta'_1 s = 0.$$

Therefore,

$$B_{m,Q} = B_{m-1,Q} \mathcal{K}_{QQ} - (B'_{m-1,Z} \Omega_{ZZ} + B_{m-1,Q} \Omega_{QZ}) \Lambda_{1,ZQ}.$$

Rewriting this expression for  $m = n - h + 1$ , we obtain from Equation (16) the expression for the impact factors

$$\gamma_h^n = -\frac{1}{n} (B'_{n-h,Z} \Omega_{ZZ} + B_{n-h,Q} \Omega_{QZ}) \Lambda_{1,ZQ}. \quad (23)$$

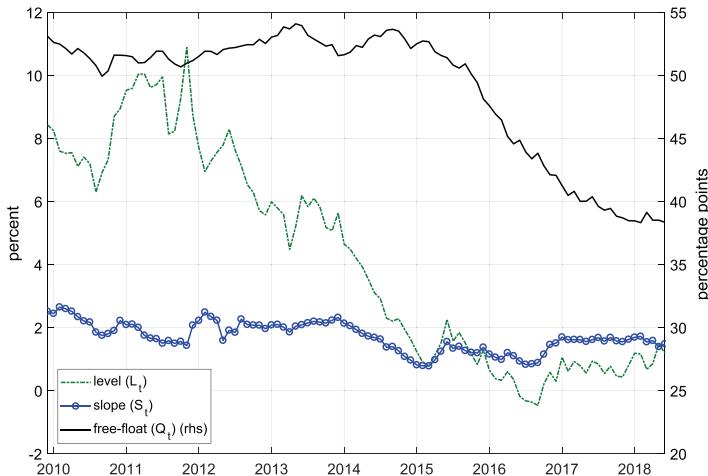
This is the same expression as in the last line of Equation (21). Therefore, an anticipated innovation  $u_{t+h-1}$  to the free-float has the same impact on the term premium as a change in the expected future free-float  $E(Q_{t+h-1}|X_t)$  due to a change in current  $Q_t$ .

Overall, the model used in this paper is inheriting key features from the equilibrium model introduced by Greenwood and Vayanos (2014) and Vayanos and Vila (2021). A more detailed comparison between the empirical model and certain properties of their theoretical framework is provided in the appendix.

## 4. Estimation

### 4.1 Estimation Approach

While we rely on the same modeling framework as Li and Wei (2013), we modify their two-step estimation approach in order to address specific challenges posed by the euro area data. In the first step we estimate a VAR of the risk factors, including the free-float variable, and the relation between the short-term rate and these factors. In the second step, we quantify the market prices of risk by using a dual objective: we simultaneously match the time-series evolution of bond yields between December 2009 and August 2014, as well as the portion of the yield curve decline between September 2014 and March 2015 that can be attributed to markets gradually pricing in expectations for large-scale asset purchases by the Eurosystem.

**Figure 6. Model Factors**

**Note:** The figure shows end-of-month time series of the level and the slope factors (left axis) and the free-float factor expressed as a percentage of total duration supply (right axis).

Specifically, in the first step we fit a VAR(1) to an empirical level and slope factor ( $L_t$  and  $S_t$ , respectively) and to our observed free-float measure  $Q_t$  (see Equation (1)), over the pre-APP subperiod from December 2009 to August 2014. The level and the slope factors are extracted as the first two principal components from the cross-section of observed yields with maturities 1-year, 2-year, ..., 10-year. Figure 6 shows monthly time series of the three factors over the full sample from December 2009 to June 2018. The shock impact matrix  $\Omega$  is the Cholesky decomposition of the variance-covariance matrix of the reduced-form shocks implied by the estimated VAR model. We estimate the parameters  $\delta_0$  and  $\delta_1$  in Equation (2) with OLS. For the VAR and the OLS regression we impose the zero restrictions on  $\mathcal{K}$  from Equation (4) and  $\delta_1$  from Equation (2), respectively.

In the second step, we estimate the market-price-of-risk parameters. In theory, we could follow Li and Wei (2013) and match the observed time series of bond yields and term premia estimates obtained from an auxiliary term structure model (Kim and Wright 2005) that excludes bond supply information. However, in practice two aspects of the euro area data prevent us from relying on such a

pure time-series approach. First, our sample is relatively short due to the limited availability of the euro area free-float measure, which is available only from December 2009. Second, Eurosystem bond holdings only became a sizable source of variation in the free-float with the start of APP. By contrast, the Federal Reserve's SOMA (System Open Market Account) portfolio exhibited significant variations already before the inception of the Federal Reserve's LSAPs. Hence, based on the euro area data, it is more challenging for the model to learn about the parameters from the covariation of  $Q$  and bond yield dynamics.

Therefore, for estimating the market-price-of-risk parameters, our second step uses a dual objective function that not only takes into account the time-series fit of bond yields but also the model's ability to capture the initial decline of the yield curve from September 2014 to March 2015. From September 2014, expectations about a possible ECB future large-scale asset purchase program were building in financial markets ahead of the start of the APP in March 2015.

For the first part of the objective function, denote by  $y_t^o$  the cross-section of observed yields with maturities 1-year, 2-year,  $\dots$ , 10-year, and by  $\hat{y}_t \equiv \hat{y}_t(\lambda_0, \Lambda_1 | X_t; \hat{c}, \hat{\mathcal{K}}, \hat{\Omega}, \hat{\delta}_0, \hat{\delta}_1)$  the corresponding fitted yields, using Equation (9) and taking the estimated VAR and short-rate parameters from the first step as given. Our distance measure is the average (across maturities and time) squared fitting error using end-of-month yields from December 2009 to August 2014:

$$F_1(\lambda_0, \Lambda_1) = \frac{1}{M_1 T} \sum_{t=1}^T [y_t^o - \hat{y}_t]' [y_t^o - \hat{y}_t], \quad (24)$$

where  $T = 56$  denotes the number of time-series observations and  $M_1 = 10$  the number of maturities used in the cross-section.

For the second part of the objective function, we assume that part of the euro area bond yield decline observed from September 2014 to March 2015 (when the APP was officially launched) was due to the buildup of private sector expectations of large-scale sovereign bonds purchases. To infer the cumulative yield decline over this time window that can be attributed to the anticipation of the APP, we conduct an event study. We then match the observed APP-induced cumulative change in bond yields with the model-implied

change in term premia, conditional on a proxy for the prevailing APP expectations at the time.

For the selection of events with APP-related news, we follow Altavilla, Carboni, and Motto (2021) and focus on a set of event dates at which the ECB conveyed news about the APP in the form of ECB press conferences as well as speeches given by ECB President Draghi. The first date is September 4, 2014, the day of the ECB press conference at which the initial purchases under the ABSPP and the CBPP3, which preceded the announcement of the APP in January 2015, were communicated. Moreover, at the same point, President Draghi indicated that a “broad asset purchase programme was discussed, and some Governors made clear that they would like to do more.”<sup>23</sup> The last date is March 5, 2015, when the ECB announced final technical details of the program, which complemented the information provided at the press conference following the January 22, 2015 Governing Council, and which confirmed March 9, 2015 as the starting date for the APP.<sup>24</sup>

For our event study we analyze the changes of zero-coupon yields over two-day windows.<sup>25</sup> We assume that the observed changes in yields around those event dates are primarily driven by market participants’ changing expectations about the APP. Following Altavilla, Carboni, and Motto (2021), we conduct two versions of the event study: one in which we control for news about key macroeconomic variables on those event dates, and another without such controls. We restrict our focus on the medium- and long-term segment of the yield curve and disregard changes of yields with less than five-year maturity. This is motivated by the fact that average short-rate expectations over shorter horizons may also reflect monetary policy news unrelated to the APP. From September 3, 2014

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<sup>23</sup>See <https://www.ecb.europa.eu/press/pressconf/2014/html/is140904.en.html>.

<sup>24</sup>Our selection of relevant dates is aligned with the working version of Altavilla, Carboni, and Motto (2021) where March 9, 2015—the date when actual APP purchases started—is not included in the set of events.

<sup>25</sup>Most of the relevant ECB announcements were made in the afternoon on a given day. We consider two-day rather than one-day yield changes, as the construction of zero-coupon yields (see Section 2.3) for a given day may incorporate prices prevailing before noon. Thus, if the announcement took place at date  $t$ , some of the price changes underlying the construction of zero-coupon yields between  $t - 1$  and  $t$  may not reflect the event of day  $t$ ’s afternoon.

to March 6, 2015 zero-coupon bond yields declined by 89 bps at the 10-year maturity (and 46 bps at the 5-year maturity). Averaging the results of the two event-study analyses (controlled versus uncontrolled), we attribute cumulative reductions of the 10-year (and 5-year) zero-coupon bond yield of 48 bps (and 33 bps) to APP announcements.

To operationalize the second component of the objective function, let  $dy^o$  denote the change in bond yields for maturities 5-year, 6-year, ..., 10-year, which is attributable to news about the APP as estimated by the aforementioned event-study approach. For example,  $dy_{10y}^o = -48$  bps. Let  $\hat{dy} \equiv \hat{dy}(\lambda_0, \Lambda_1 | \mathcal{U}_t; \hat{\epsilon}, \hat{\mathcal{K}}, \hat{\Omega}, \hat{\delta}_0, \hat{\delta}_1)$  denote the corresponding model-implied changes over the same period, computed by deploying Equation (16) for the respective maturities. The  $\mathcal{U}$  sequence used in (16) represents the expected trajectory for duration extraction determined by the APP as of March 5, 2015. This trajectory is constructed based on survey expectations prevailing at that date, which were closely aligned to the Governing Council's January 2015 announcements. We assume that this  $\mathcal{U}$  sequence represents the APP expectations prevailing when the program was launched, as they had built up "from zero" from September 2014. The second part of the objective function is then given by the distance measure:

$$F_2(\lambda_0, \Lambda_1) = \frac{1}{M_2} [dy^o - \hat{dy}]' [dy^o - \hat{dy}], \quad (25)$$

where  $M_2 = 6$  denotes the number of maturities used in the second part of the objective function.

The optimization problem for estimating the market-prices-of-risk parameters then is

$$\{\hat{\lambda}_0, \hat{\Lambda}_1\} = \arg \min_{\{\lambda_0, \Lambda_1\}} \omega F_1(\lambda_0, \Lambda_1) + (1 - \omega) F_2(\lambda_0, \Lambda_1), \quad (26)$$

where  $\omega$  is a weighting parameter that balances the importance of the time-series fit criterion  $F_1$  and the "event-window" fit criterion  $F_2$  for the overall objective function. Choosing  $\omega$  requires judgment. With a view of imposing a "flat prior" across the two criteria, we set  $\omega = 0.5$ . However, altering the weight  $\omega$  does not affect our estimates of parameters and yield curve impacts very much, as it turns out that

the events component mainly informs the market-price-of-risk parameters governing the mapping from free-float to term premia that are of relatively minor relevance for the overall time-series fit over the first part of our sample.

#### *4.2 Parameter Estimates*

Table 3 reports estimates of the model parameters. Using estimates of the price-of-risk parameters  $\lambda_0$  and  $\Lambda_1$ , which are derived in the second estimation step, we compute estimates for the parameters  $\tilde{c}$  and  $\tilde{\mathcal{K}}$  that govern the risk-neutral dynamics of factors  $X_t$ . The higher eigenvalues of  $\tilde{\mathcal{K}}$  than  $\mathcal{K}$  indicate that all three factors are more persistent under the risk-neutral ( $\mathbb{Q}$ ) than the real-world ( $\mathbb{P}$ ) probability measure.

Given these parameter estimates and using the affine relation (9) between factors  $X_t$  and bond yields  $y_t^n$ , we report in Table 4 the reaction of the yield curve to a time- $t$  increase in each factor, which amounts to one standard deviation of the reduced-form shocks derived from the estimate of the shock variance-covariance matrix  $\Omega\Omega'$ . The first column of Table 4 indicates that the loadings of yields on the level factor  $L_t$  are positive and of similar size across maturities. Thus, a positive shock to this risk factor leads to an (almost) parallel upward shift of the entire yield curve. A positive shock to the slope risk factor  $S_t$  leads to a steepening of the yield curve, as indicated by the second column of Table 4.

A contemporaneous shock to the quantity factor  $Q_t$  shifts the entire yield curve in the same direction as the shock, as indicated by the third column of Table 4. Therefore, in line with economic intuition, yields decrease when the free-float measure is reduced by the Eurosystem's duration extraction. The yield impact of a shock to the free-float is hump-shaped across maturities. This is one of two possible shapes that can arise in the equilibrium model by Greenwood and Vayanos (2014). As argued by the authors, the hump-shaped pattern can occur when the shock to the free-float is mean-reverting relatively quickly. Indeed, according to the estimated  $\mathcal{K}_{QQ}$  of 0.9039 from Table 3—which represents the persistency of the  $Q$  factor due to the imposed restrictions on the interactions with the other two factors—the impact of a shock to the free-float has a half-life of only seven months. In addition, we find that a shock to the

**Table 3. Parameter Estimates**

	$\delta_0$ 0.0045			
	$\delta_1$			
$L_t$ $S_t$ $Q_t$	0.1240 -0.4432 0			
		$\mathcal{K}$		$c$
$L_t$ $S_t$ $Q_t$ eig ( $\mathcal{K}$ )	0.9835 0.0150 0 0.8256	-0.0748 0.8185 0 0.9039	0 0 0.9039 0.9764	0.0016 0.0024 0.0501
		$\Omega$		
$L_t$ $S_t$ $Q_t$	0.0067 -0.0003 0.0001	0 0.0018 0.0002	0 0 0.0041	
		$\Lambda_1$		$\lambda_0$
$L_t$ $S_t$ $Q_t$	-5.1457 33.4122 0	-17.5141 -63.7912 0	-0.5450 1.4627 0	0.5768 -1.2922 0
		$\tilde{\mathcal{K}}$		$\tilde{c}$
$L_t$ $S_t$ $Q_t$ eig ( $\tilde{\mathcal{K}}$ )	1.0181 -0.0450 -0.0062 0.9044	0.0429 0.9253 0.0132 0.9567	0.0037 -0.0027 0.9036 0.9860	-0.0023 0.0048 0.0504

**Note:** This table reports parameter estimates of the model obtained by using the two-step approach described in Section 4.1. In the first step we derive estimates for  $c$ ,  $\mathcal{K}$ , and  $\Omega$ —which govern the real-world dynamics of factors  $X_t$ —and of  $\delta_0$  and  $\delta_1$ —which map linearly factors  $X_t$  into the short rate—and, in the second step, for market-price-of-risk parameters  $\lambda_0$  and  $\Lambda_1$ . Given these estimates, we compute  $\tilde{c} = c - \Omega\lambda_0$  and  $\tilde{\mathcal{K}} = \mathcal{K} - \Omega\Lambda_1$ , which govern the risk-neutral dynamics of factors  $X_t$ .

free-float moves the contemporaneous one-period expected returns  $E_t rx_{t+1}^{n-1} = B'_{n-1} \Omega \lambda_t$  of all bonds in the same direction as the shock, and that the effect is increasing across maturities. Also this empirical finding of our euro area model is in line with the prediction of the

**Table 4. Reaction of the Yield Curve  
to Changes in the Factors**

Maturity of Yield (Years)	$L_t$	$S_t$	$Q_t$
1	15	-5	0.29
2	20	-2	0.41
3	22	-1	0.44
4	23	0	0.44
5	23	1	0.42
6	23	1	0.41
7	23	2	0.38
8	22	2	0.36
9	21	2	0.34
10	20	2	0.32

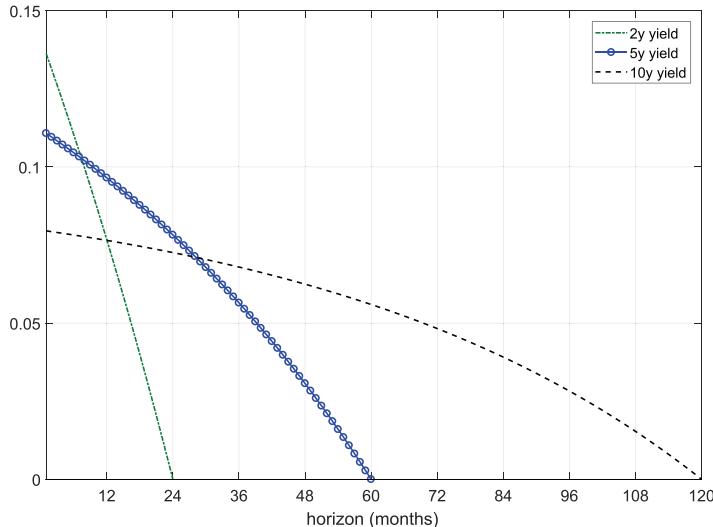
**Note:** The table reports changes in yields in response to a one-standard-deviation shock in time- $t$  to each factor in  $X_t = (L_t, S_t, Q_t)'$ . The changes are derived using Equation (9). The changes are reported in basis points.

theoretical model of Greenwood and Vayanos (2014); see Section 1.3 of their paper.

Finally, we inspect the estimated impact factors  $\gamma$  that map sequences of innovations of the free-float from the expected path implied by the VAR dynamics—Equation (3)—into yield curve reactions. Figure 7 plots the estimated impact factors  $\gamma_h^n$  for maturities  $n = 2, 5$ , and 10 years over the relevant horizons  $h$ . The impact factors decrease monotonically over the future horizons within the tenor of any bond. Therefore, changes in free-floating duration supply over the near term have a larger effect on the term premium component of a yield than changes occurring in the more distant future. This pattern holds for bonds of any maturity.

#### 4.3 Model Fit

For our time-series fitting criterion  $F_1$  in (26), the model delivers a good fit to the yield-level data over time. Root mean squared errors over the estimation period range from 3 to 14 bps across maturities. This is comparable in size to the fit of U.S. yields in Li and Wei

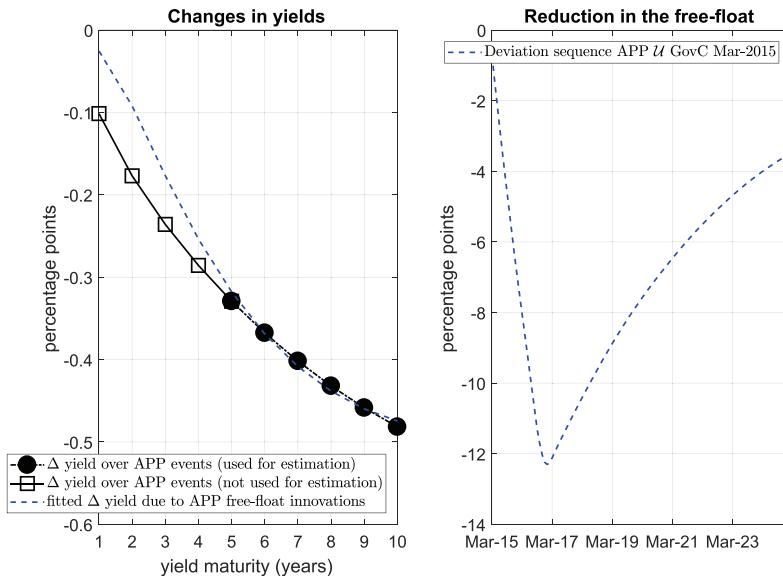
**Figure 7. Estimated Impact Factors**

**Note:** The figure shows the estimated impact factors  $\gamma_h^n$  that map a sequence of revisions to the current and expected free-float into changes in yields of bonds with maturity  $n = 2, 5$ , and  $10$  years; see Equation (16). The vertical axis shows the yield change contribution per unit of free-float change at the respective horizon. For instance, an anticipated free-float reduction in three years by 1 percentage point contributes to lowering the five-year yield by 0.05 percentage point contemporaneously.

(2013). In particular, medium- and long-term yields are fitted well in sample.

Focusing on  $F_2$ , the second part of the objective function in (26), the left panel of Figure 8 presents the model fit of the APP-induced cumulative decrease in the sovereign yield curve over selected event dates from September 2014 to March 2015, as discussed in Section 4.1. The right panel of Figure 8 plots the series of anticipated future APP-induced free-float innovations  $\mathcal{U}$  as of March 5, 2015 that underlies these fitted yield changes. Each model-implied yield decline shown is the result of multiplying the maturity-specific impact factors  $\gamma_h^n$ —see Equation (16)—with the part of this  $\mathcal{U}$  sequence spanning the horizons of each yield tenor. For example, the 10-year yield APP-induced compression (of about 49 bps) is the product of the  $\gamma_h^{10y}$ , the dashed line in Figure 7, and the part of the  $\mathcal{U}$  sequence

**Figure 8. Impact of the Anticipation of the APP on the Yield Curve through Expected Future Free-Float Compression**



**Note:** The left panel plots observed and fitted changes in yields over an event window; the black line represents the cumulative decreases in yields over the APP-related events from early September 2014 to early March 2015; the blue dashed line plots the model-implied changes in yields due to future APP-induced free-float innovations  $\mathcal{U}$  as of March 5, 2015 (shown in the right panel). Decreases in yields with maturities as of 5 years up to 10 years (circle markers) are used in the estimation of the model, while the decreases of yields with shorter maturities (square markers) are left out.

from the right panel of Figure 8 starting in March 2015 and ending in February 2024.

The model fits almost perfectly the decreases in yields with maturities of five years and more, which corresponds to the data used in the second part of the objective function. For shorter maturities, which do not enter the estimation criterion, the model predicts less pronounced yield decreases than observed for the selected events. As the model captures only the effect of the APP on term premia due to the duration extraction channel, the observed undershooting of the model prediction is attributable to factors outside our model, such

as a signaling channel of APP-related communication, or changes in the expected ECB's key interest rate policy rates.

The reduction in the future free-float induced by the anticipation of the APP—see the right panel of Figure 8—that underlies the fitted decreases in yields between September 2014 to March 2015 is large relative to the average variation of the supply factor from December 2009 to August 2014. The standard deviation of innovations to the free-float in this early period amounts to only 0.4 percentage point (see the estimate of the shock variance-covariance matrix  $\Omega\Omega'$  from Table 3). By contrast, in March 2015 the anticipated reduction in the free-float induced by the APP was envisaged to peak at about 12 percentage points at the end of the net purchases phase and to still amount to about 4 percentage points in 2025; see again Figure 8. Hence, in contrast to the short-lived persistence of a shock to the free-float in the pre-APP period (with a half-life of only seven months; see Table 3), the APP represents a very persistent reduction in the supply of available bonds. In terms of implementation, recall that we condition in the bond pricing equations on arbitrary free-float processes—in turn based on ECB policy announcements and market participants' expectations gleaned from surveys—thereby “overwriting” the free-float dynamics implied by the estimated VAR. In the standard model setup without such feature, any free-float-induced change in yields would need to come via a change in the contemporaneous state variable, i.e., via an adequately sized innovation, and expectations of future free-float would follow from VAR dynamics. Here, following Li and Wei (2013), we can directly condition on a given free-float path; see again Equation (13). To illustrate further the extraordinary dimension of the APP compared with historical free-float variation, we can compute the unanticipated contemporaneous free-float shock that, when multiplied with the respective yield loading  $B_{10y,Q}$ , would give the same 49 bps impact on the 10-year yield as the anticipated free-float shock sequence as of March 2015. It turns out that this hypothetical shock would amount to 61 percentage points. This represents more than the actual supply of bonds available to arbitrageurs at the time the program was launched; see Figure 6. Overall, both the size and the persistence of APP-induced innovations to future free-float are several orders of magnitude higher than the free-float variation observed in the pre-APP sample.

## 5. The Impact of the APP on the Yield Curve

We use our estimated model to infer the impact of the APP on the sovereign yield curve through the duration extraction channel. First, we estimate the compression of term premia along the yield curve for different vintages of the APP (Section 5.1). Second, we examine the persistence of the term premium compression over time and investigate the contribution made by reinvestments of maturing principals (Section 5.2). Third, we compare the yield changes observed around APP recalibration announcements to the real-time predictions of our model (Section 5.3). Finally, we assess the robustness of our results (Section 5.4).

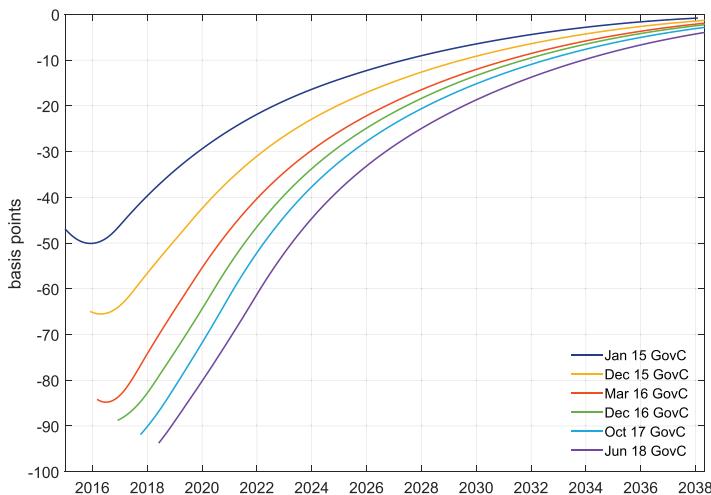
### 5.1 Term Premium Compression across the Yield Curve

Figure 1 shows the estimated impact of the APP across the yield curve for the different APP vintages at the time of their announcement. Each curve shows the estimated term premium compression relative to the counterfactual of no duration extraction through the APP. The date shown in the legend indicates both the respective APP vintage, i.e., the specific path of net purchases implied by the vintage, as well as the point in time at which the term premium compression is estimated. For example, the curve labeled “Jan 15 GovC” shows the estimated term premium compression due to the January 2015 APP vintage in January 2015.

We obtain these yield curve impact estimates by feeding the free-float reduction implied by the different APP vintages (see Figure 4) into our model. In detail, the estimated term premium compression is constructed using Equation (16), which maps the sequence of anticipated free-float innovations ( $\mathcal{U}$ ) into a yield impact using the impact factors (see illustrations of  $\gamma_h^n$  for selected maturities in Figure 7). For the example of January 2015, the sequence of free-float innovations relevant for the 10-year term premium is the part of the dark-blue line corresponding to the January 2015 APP vintage in Figure 4 that ranges from January 2015 10 years to December 2024. To compute the five-year term premium, the relevant sequence of free-float innovations consists of just the five years from January 2015 until December 2019 in Figure 4.

For January 2015, the impact on the 5-year and 10-year term premium is found to be around 30 and 50 bps, respectively. Also for

**Figure 9. The Impact of the APP on the 10-Year Term Premium over Time**



**Note:** For selected dates, the figure shows the estimated contemporaneous and future impact of the APP on the term premium component of the 10-year sovereign bond yield.

the subsequent vintages, the term premium impact is estimated to be higher for longer tenors, i.e., the APP has led to a flattening of the curve. The overall term structure impact has become stronger over time as the APP has been expanded in length and volume. For the June 2018 APP vintage we estimate that in the absence of the APP the 10-year sovereign bond yield would have been around 95 bps higher at that point (Figure 1).

## 5.2 Term Premium Compression over Time

Figure 9 plots the term premium impact at the 10-year maturity for different APP vintages over time. At each point in time indicated on the horizontal axis, the figure shows the estimated 10-year term premium compression for the different APP vintages.

Figure 9 is constructed as follows. For each of the trajectories shown there, the starting point is the initial impact, i.e., the 10-year maturity point in Figure 1. For each of the trajectories shown, the impact over time is then obtained by moving to the right along the

corresponding free-float compression curve in Figure 4. To this end, we use the impact formula (16)—reproduced here for convenience:  $dy_n(\mathcal{U}_{t+h}) = \gamma^{n'}\mathcal{U}_{t+h}$ —by applying the impact vector  $\gamma^{10y}$  to the sequences of anticipated innovations  $\mathcal{U}$  that start in the future at time  $t + h$ . For example, for the June 2018 vintage we estimate the 10-year term premium reduction in January 2025, by taking the segment of the violet free-float impact curve in Figure 4 that starts in January 2025 and ends in December 2034 as our sequence for the free-float reduction. The scalar product with the time-invariant impact factor vector ( $\gamma^{10y}$ ) then delivers an estimated 10-year yield impact of around 35 bps in January 2025.

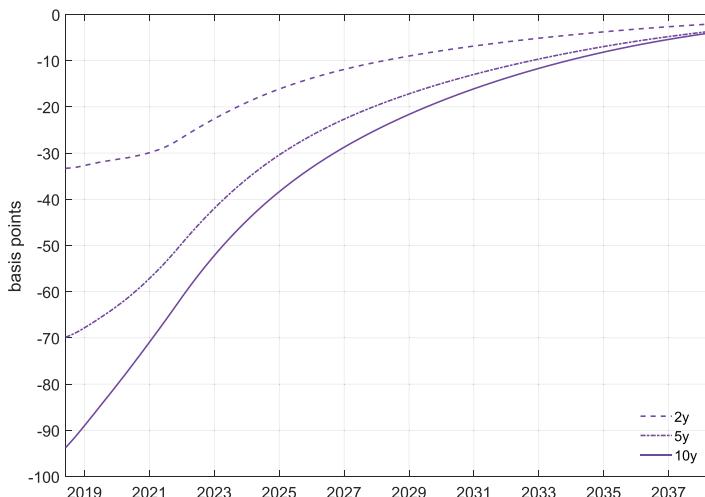
The estimated term premium impact is fairly persistent but gradually fades over time. Across the APP trajectories shown, the half-life of the initial impact on the 10-year yield is around five to six years. While the projected 10-year term premium compression falls below 10 bps by around 2033, it only dissipates completely once the portfolio has been entirely wound down.

For shorter maturities, the impact of the program also diminishes over time, albeit more slowly than at longer maturities; see Figure 10 for the June 2018 APP vintage. Looking at the 2-year maturity, the initial term premium effect is smaller than for the 10-year maturity, which implies a flattening of the curve, as discussed above. As the 2-year impact fades more slowly than the 10-year impact, the yield curve becomes again steeper over time. The markedly greater persistence of the two-year term premium compression over the nearer term reflects the impact of reinvestments, which were anticipated to follow the end of net purchases in December 2018 and assumed to last for three years. Hence, even in early 2020 most of the term of a two-year bond is falling into the reinvestment phase, which is not true for longer-term bonds.

The fading of the term premium compression reflects, to some extent, the “aging” of the portfolio—i.e., its gradual loss of duration as the securities held in the portfolio mature—as well as, in particular, the run-down of the portfolio that market participants anticipate will eventually follow the end of the expected horizon of reinvestments.

The pure “aging” effect is due to the fact that day by day the duration of the central bank portfolio falls even in the absence of any redemptions. The reinvestment of maturing principals conducted in

**Figure 10. The Impact of the APP on the 2-Year, 5-Year, and 10-Year Term Premium over Time**



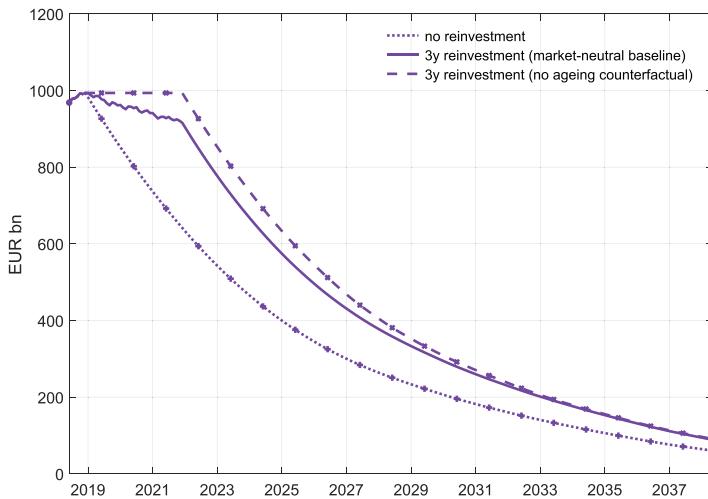
**Note:** The figure shows for the June 2018 APP vintage the impact of the APP on the term premium component of the 2-year, 5-year, and 10-year sovereign bond yield (averaged across the four largest euro area countries) over time.

line with “market neutrality”—i.e., with the maturity distribution of purchases guided by the maturity distribution of the eligible universe of securities—offsets this gradual loss of duration to some extent (see the continuous versus the dotted line in Figure 11) over the reinvestment horizon (assumed to be two years). By contrast, under a counterfactual “no aging” reinvestment policy (dashed-dotted line in Figure 11), the portfolio would remain constant in terms of 10-year equivalents during the reinvestment phase. Figure 12 illustrates the term premium compression that would result from such a counterfactual “no aging” reinvestment policy.<sup>26</sup> It turns out that even if the portfolio was prevented from aging during the reinvestment phase, the term premium impact of the central bank purchases would still fade gradually over time. This suggests that the bulk of the fading

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<sup>26</sup>This type of reinvestment policy would be challenging to implement in practice, as it would require reinvestments into very long-term securities, with an average maturity of around 13 years.

**Figure 11. Illustrating Portfolio Aging: The Evolution of Duration-Weighted Government Bond Holdings Under Different Reinvestment Scenarios**

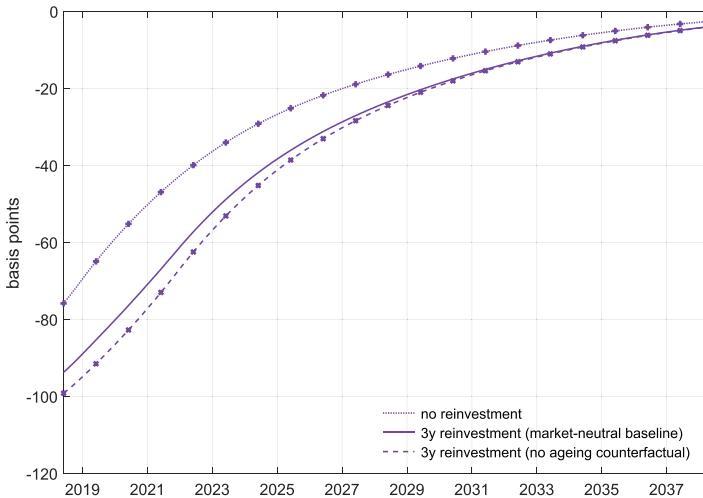


**Note:** For the June 2018 APP vintage of net purchases, the figure shows the projected evolution of the big-four government bond holdings in terms of 10-year equivalents under three reinvestment scenarios. Under the “no reinvestment” scenario, the portfolio starts running down after the end of net purchases in December 2018. In the “3y reinvestment scenario (market-neutral baseline)” scenario, reinvestments are made for three years starting in January 2019 in line with a “market-neutral” maturity distribution of purchases. In the “3y reinvestment (no ageing counterfactual)” scenario, reinvestments are again conducted over a period of three years starting in January 2019, but deviating from our baseline case it is assumed that reinvestments are made in sufficiently long maturities to offset the “aging” of the portfolio during the reinvestment phase.

term premium impact in the future reflects market expectations of a gradual roll-down of the portfolio after the end of reinvestments.

Apart from the relevance of reinvestment in mitigating the aging effect, the reinvestment horizon as such makes an important contribution to the reduction in term premia and its persistence over time. Figure 13 illustrates for the June 2018 APP vintage the 10-year term premium reduction for reinvestment horizons ranging from 0 to 10 years. The longer is the reinvestment horizon, the higher is the contemporaneous yield impact. However, the marginal impact of an additional year of reinvestment is shrinking with the length of the reinvestment horizon. For instance, reinvesting for 3 years instead

**Figure 12. Illustrating Portfolio Aging: The APP's 10-Year Term Premium Impact under Different Reinvestment Scenarios**



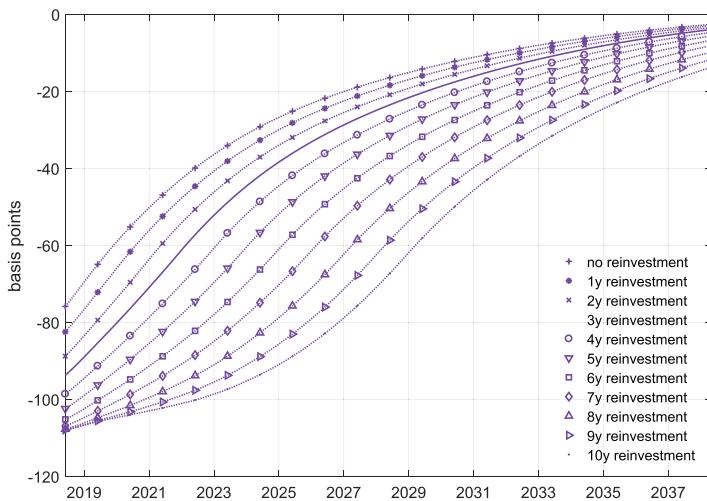
**Note:** The figure shows the 10-year term premium impact over time that is implied by the respective trajectory of central bank holdings shown in Figure 11.

of 0 years generates an additional term premium impact of around 18 bps, while moving from 7 to 10 years of reinvestment induces an additional compression of a mere 2 bps. This declining marginal effect reflects discounting: the additional free-float reduction in the 7-year versus 10-year reinvestment scenario happens 7 years from now, which is priced into contemporaneous term premia via low levels of impact factors; see Figure 7 again. But the picture changes over time: standing in, say, 2026, the marginal impact of going from 7-year to 10-year reinvestment (following the end of net purchases in December 2018) is larger than in June 2018.

### 5.3 Benchmarking Announcement Effects of APP Recalibrations

We benchmark the term premium impact of APP recalibration announcements predicted by our model against the observed yield curve changes within narrow windows around the respective announcement dates. Since our model estimation is not informed

**Figure 13. Evolution of the 10-Year Term Premium Compression through the APP for Different Reinvestment Horizons**



**Note:** For the June 2018 APP vintage, the figure shows the 10-year term premium compression over time for different reinvestment horizons following the end of net asset purchases in December 2018.

by post-March 2015 data, these exercises represent an out-of-sample cross-check of our model.

To calculate the surprise entailed by the APP recalibration announcements for the future free-float, we control for pre-announcement market expectations. For each APP recalibration we first simulate the free-float trajectory based on survey expectations about the path of the APP before the announcement and then again based on the actually announced purchase parameters (see Table 2). The difference between these two free-float trajectories gives us the sequence of surprises to the free-float due to the APP recalibration announcement. We feed these surprises into our model (using them as the  $\mathcal{U}$  sequence in Equation (16)), and compare this model prediction with the one-day yield curve changes measured around the APP recalibration announcement date.

We cleanse the observed yield curve changes from both changes in the bond yield's expectations component (average short-rate expectations over the bond's maturity) as well as macro surprises.

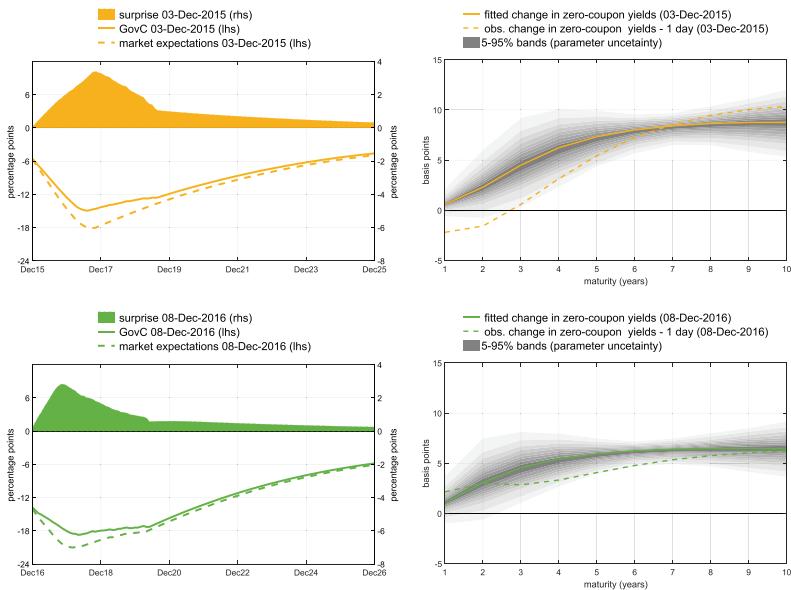
This makes the yield changes more closely comparable to the yield changes predicted by our term structure model, which captures the change in yields purely based on the term premium compression via duration risk extraction. Specifically, first, to control for the expectations component we subtract from the full observed yield change the change in the estimated expectation component of the euro area swap (OIS) rate curve, which we obtain from a benchmark affine term structure model based on Joslin, Singleton, and Zhu (2011). Second, we account for macroeconomic surprises on the days of the announcements of APP recalibrations by cleansing yield changes for macro effects relying on the sensitivity of yields to macroeconomic surprises obtained in Section 4.1. The yield changes shown on the right-hand side of Figure 14 are the average of the observed yield changes cleansed of macro effects and those not cleansed of such effects (but in both cases cleansed of changes in short-rate expectations).

For the December 2015 and December 2016 APP recalibrations, Figure 14 shows on the left side the surprises in the free-float sequences and on the right side the corresponding model-implied changes in the yield curve, as well as the (cleansed) observed yield curve changes. Among the dates at which the ECB recalibrated its purchase program, the December 2015 and December 2016 announcements stand out as those that surprised the market the most and triggered the strongest yield response. Both recalibration announcements implied less duration extraction by the ECB than the market anticipated and therefore some upward revisions to the free-float of duration risk.<sup>27</sup> Yields increased for all but the

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<sup>27</sup>In December 2015, the ECB announced the first recalibration of the APP since the initial announcement of the program in January 2015. The recalibration involved the announcement of, first, a prolongation of net purchases by six months until March 2017 at an unchanged purchase pace of €60 billion per month, and, second, a reinvestment policy for maturing principals beyond the net purchase horizon “for as long as necessary.” While market participants had anticipated the prolongation of net asset purchases ahead of the December 2015 Governing Council, the predominant expectation had been for the ECB to also increase the monthly pace of purchases (see Table 2). As a result, and despite the announcement of the reinvestment policy, the APP recalibration implied a significantly lower duration absorption over the near term than expected by market participants. The December 2016 recalibration featured an extension of net purchases at a reduced monthly pace of €60 billion for nine months until December 2017.

**Figure 14. Announcement Effects of Selected APP Recalibrations**



**Note:** Left-hand side: The dashed lines represent the expected APP-induced free-float compression paths constructed from surveys before the APP recalibration was announced. The continuous lines represent the expected free-float compression paths after the APP recalibration announcement, constructed based on the announced APP parameters. The difference between the pre- and post-APP recalibration free-float projections is the free-float surprise, which is shown as an area. The right-hand side shows the model-implied term premium impact due to the free-float surprises (APP recalibrations versus observed yield changes). The observed changes (cleansed from changes in short-rate expectations and macroeconomic surprises) are shown as dashed lines. The gray area is the bootstrap-based 5–95 percent confidence band of the model-implied yield changes reflecting parameter uncertainty.

shortest maturities on the day of the Governing Council announcement, thereby steepening the curve. Filtering the resulting revision

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Market expectations were for an extension over a somewhat shorter horizon at a slightly higher monthly pace; see Table 2. In addition, the ECB expanded the eligible maturity bracket from two to one year at the lower end and also opened the door to purchases at yields below the deposit facility rate “to the extent necessary.” The somewhat lower monthly purchase pace and increased scope for buying short-term papers implied some upward surprise on the expected path of the free-float.

of the free-float sequence through our model can broadly explain the observed reaction of the yield curve: the model predicts the right shape of change, i.e., the curve steepening, and (cleansed) observed yield changes are overall covered by our bootstrap-based 5–95 percent confidence intervals for model-implied yield changes.<sup>28</sup>

By contrast, the remaining APP recalibrations were closely in line with market expectations. As a result, for those recalibrations there was no revision to the expected free-float of duration risk that helps identify how well the model captures revisions of the free-float sequence. In the absence of such revisions, our model predicts no changes in yields for those APP recalibrations. On those days, nonetheless, some mild movements in the yield curve were observed, which likely reflect changes in other policy instruments—in particular, the cut of the deposit facility rate that accompanied the March 2016 APP recalibration and the change in the ECB’s forward guidance on the path of policy rates that came with the June 2018 APP recalibration.

#### *5.4 Uncertainty and Robustness of Results*

We conduct sensitivity analyses around our baseline results via four avenues: first, we account for parameter uncertainty based on bootstrapping; second, we conduct a bias adjustment of the estimated factor dynamics; third, we vary our estimation sample; fourth, we reestimate the model with a differently scaled bond supply variable.

To account for parameter uncertainty, we rely on a bootstrap procedure. We do so, as our two-step estimation approach and the limited number of available observations prevent a straightforward application of asymptotic results. For the bootstrap we resample the data and obtain bootstrap estimates of the model parameters based on our two-step estimation approach outlined in Section 4.1.

In detail, in the  $i$ th bootstrap run, for the first step of our estimation approach, we take random draws from the centered residuals of our estimated factor VAR and use them as innovations in generating a new time series of factors, based on the point estimates

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<sup>28</sup>See Section 5.4 for details on the construction of the confidence intervals.

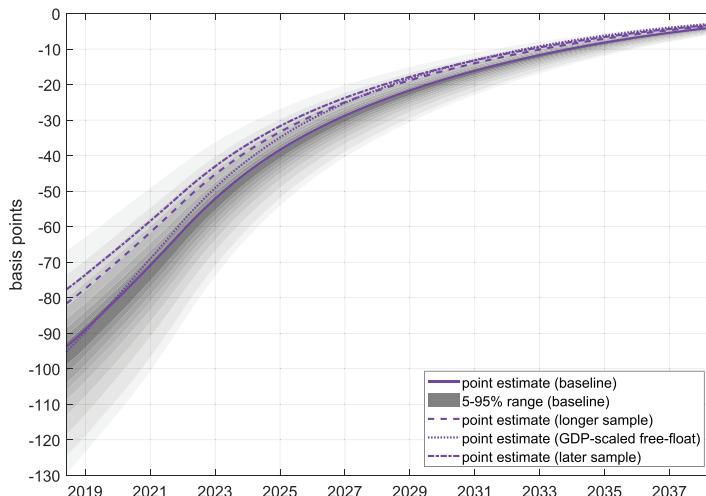
of VAR parameters  $\hat{c}$  and  $\hat{\mathcal{K}}$ . Using those bootstrap factor realizations, we reestimate the factor VAR parameters—under the same zero restrictions as in our baseline estimation—and obtain bootstrap estimates  $\check{c}^{(i)}$ ,  $\check{\mathcal{K}}^{(i)}$ , and  $\check{\Omega}^{(i)}$ . Similarly, we reestimate the short-rate equation (2) and obtain bootstrap estimates  $\check{\delta}_0^{(i)}$  and  $\check{\delta}_1^{(i)}$  of  $\delta_0$  and  $\delta_1$ . For the second step of our estimation, we bootstrap the realizations of the components featuring in our dual objective function (26). For the first part of that objective function,  $F_1$ , we construct a bootstrap realization of the time series of yields by adding measurement errors to fitted yields, where these errors are drawn from the pool of centered fitting residuals of our estimated model. For the second component,  $F_2$ , we then generate a bootstrap realization of the change in the yield curve over our event window by applying noise around the fitted yield changes.<sup>29</sup> Given the bootstrap draw of the yield changes over the event window and the bootstrap draw of the yields sequence, we conduct the second step of our estimation approach, i.e., we minimize the dual objective criterion (26) for given  $\check{c}^{(i)}$ ,  $\check{\mathcal{K}}^{(i)}$ ,  $\check{\Omega}^{(i)}$ ,  $\check{\delta}_0^{(i)}$ , and  $\check{\delta}_1^{(i)}$  and obtain bootstrap estimates  $\check{\lambda}_0^{(i)}$  and  $\check{\Lambda}_1^{(i)}$  of  $\lambda_0$  and  $\Lambda_1$ , respectively. We repeat this procedure for  $K = 1,000$  bootstrap repetitions. Collecting our parameters in a vector  $\theta$ , the distribution of our point estimate  $\hat{\theta}$  is hence approximated by the sampling distribution  $(\check{\theta}^{(1)}, \dots, \check{\theta}^{(K)})$  of our bootstrap estimates. Similarly, the distribution of (non-linear) functions of the parameters  $g(\theta)$ —like, e.g., the impact factors  $\gamma_h^n \equiv \gamma_h^n(\theta)$ —are approximated by the corresponding bootstrap sampling distribution  $g(\check{\theta}^{(1)}), \dots, g(\check{\theta}^{(K)})$ . This enables us to generate distributions around all our impact estimates that reflect the uncertainty stemming from parameter estimation.

Figure 15 shows the APP's dynamic impact on the 10-year term premium based on the June 2018 APP vintage with our estimated confidence bands. The midpoint (solid violet line) is the same as in Figure 9. The uncertainty band is the bootstrap-based confidence band reflecting parameter uncertainty. For the contemporaneous

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<sup>29</sup> Proceeding in the standard fashion as we did for the bootstrap generation of factors and yields, we could draw from the residuals corresponding to the fit of our six yield changes over the event window. However, as these residuals are very small (not exceeding 2 bps), we see a risk of underestimating the true uncertainty and take a more conservative approach by drawing the noise from six independent normal distributions with zero mean and a standard deviation of 10 bps.

**Figure 15. The APP's 10-Year Term Premium Impact: Parameter Uncertainty and Sample Robustness**



**Note:** Conditional on June 2018 information, the figure shows the impact of the APP on the 10-year term premium over time. The solid line is the point estimate, identical to the violet line in Figure 9. The shaded area is the 5–95 percent confidence band stemming from parameter uncertainty computed using a bootstrap approach. The dotted line (“GDP-scaled free-float”) corresponds to a model specification that uses the nominal GDP of the big-four euro area countries for scaling the free-float ratio instead of overall bond supply. The dashed line (“longer sample”) is an alternative point estimate that uses the full-sample yield information (until June 2018) for estimation, whereas the baseline estimate ignores data after March 2015. The dashed-dotted (“later sample”) line corresponds to an estimation that uses data as of September 2014 only.

term premium impact as of June 2018, the 5–95 percent confidence band ranges from 65 to 130 bps around the 95 bps estimate. The width of the confidence bands accounting for parameter uncertainty is of the same order of magnitude as that reported in Ihrig et al. (2018). Over time the uncertainty band gradually narrows, as the point estimate and the uncertainty around it converges to zero. Formally, this can be seen from the fact that at any future point in time  $t + h$  the yield impact is given by the product of impact factors and a sequence of APP free-float innovations going forward,  $dy_n(\mathcal{U}_{t+h}) = \sum_{k=1}^n \gamma_k^n(\theta) u_{t+h+k-1}$ : as the innovations  $u_{t+h+k}$  eventually shrink to zero, so does the overall scalar product.

As a second robustness check, we conduct a bias adjustment of the estimated factor dynamics. As noted in the literature,<sup>30</sup> term structure models tend to underestimate the high persistence exhibited by bond yields—in particular, when the estimation sample is short. Bauer, Rudebusch, and Wu (2012) and others have, therefore, suggested to conduct a bias correction when estimating the VAR dynamics of factors. For our model and data, in fact, the estimated degree of persistence of the factor VAR is already high, with a maximum eigenvalue of 0.976 (Table 3). Nevertheless, we apply the bias correction methods suggested by Bauer, Rudebusch, and Wu (2012).<sup>31</sup> Overall, the bias correction leaves the main results regarding the APP's impact on the yield curve essentially unchanged. We attribute this to the fact that, despite different dynamics of factors under the  $\mathbb{P}$  measure, the cross-section information in bond yields, especially their change over the event-window dates, ensures that key objects like the impact factors  $\gamma_h^n$  in (16), which depend on both  $\mathbb{P}$  and market-price-of-risk parameters, are hardly affected.

Thirdly, we vary the estimation sample. In our baseline specification we only use data up to March 2015 to estimate the model. This approach allows undertaking a clean out-of-sample benchmarking exercise over the period of the Eurosystem asset purchases, as discussed in Section 5.3. One robustness check we conducted is to use the full data set spanning December 2009 to June 2018 to estimate the model. Specifically, for the first step of the estimation approach we estimate the factor VAR and the short-rate equation over the full sample. For the second step, we leave the second component of the dual objective function in Equation (26) unchanged, but include the full time series of bond yields until June 2018 in the first component of the dual objective function. As shown in Figure 15, dashed line, using the full sample, the estimated impact of the APP is of a similar magnitude, if somewhat smaller (81 versus 93 bps on impact in June 2018; and 42 versus 48 bps after five years). Furthermore, we conducted an additional robustness check by estimating the model using only data as of the time asset purchases started being anticipated,

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<sup>30</sup>See Bauer, Rudebusch, and Wu (2012) and Kim and Orphanides (2012).

<sup>31</sup>We deploy an analytical bias approximation, a bootstrap-based bias correction, and an indirect inference estimator for bias correction, all based on the code for Bauer, Rudebusch, and Wu (2012) provided on Cynthia Wu's website.

i.e., September 2014, until June 2018. In order to ensure comparability with the baseline estimation, we left the data used for the event-study featuring in the first part of the dual objective function—September 2014 to March 2015—unchanged. The strength of the yield impact estimates (dashed-dotted line) decreases further under this setup, but only marginally compared with the one using the full data set, and it remains within the range of the bootstrap confidence bands associated with the baseline approach.

As a final robustness check, we rerun the model estimation and selected policy analysis with a different free-float measure that uses the total nominal GDP of the big-four euro area countries as scaling variable in the denominator instead of total bond supply as in our baseline specification. Doing so, we estimate a broadly similar yield impact (see the dotted line in Figure 15). This reflects the fact that the two free-float measures are highly correlated. Moreover, the reestimated parameters obtained with the alternative free-float measures partially compensate for the differences in the free-float measures.

## 6. Conclusion

Central bank bond purchases extract duration risk which otherwise would be borne by arbitrageurs. This decreases the market price of risk and compresses the term premium component of bond yields across the term structure. Our paper quantifies the strength of the duration channel for the European Central Bank’s asset purchase program (APP). We deploy an arbitrage-free term structure model along the lines of Li and Wei (2013). In addition to the level and slope factors, aggregate duration affects the market price of level and slope risk and hence the term premium across maturities. This link between bond supply and yields reflects one of the key features of the microfounded equilibrium model of Vayanos and Vila (2021).

We find that, first, the contemporaneous impact of the ECB’s APP flattens the yield curve and amounts to around 95 bps by mid-2018 for the 10-year maturity. This impact is comparable to point estimates found for the Federal Reserve’s large-scale asset purchase programs. Second, the effect is persistent and expected to only slowly fade over time, with a half-life of around five years. Third, the expected length of the reinvestment period after net purchases has a significant impact on term premia. For example, as of June

2018, relative to a counterfactual of no reinvestment, an expected reinvestment horizon of three years compressed term premia by an additional 18 bps. Finally, recalibrations of APP purchase parameters imply surprises for the central bank's expected path of duration extraction. Overall, our model accounts well—in real time—for the duration-implied yield curve impact of such recalibrations on the term structure of interest rates, while at the same time other factors—going beyond the duration channel—can move the yield curve around such announcements but are outside the scope of our model.

Our contribution to the literature is threefold. First and foremost, our paper is the first to provide a comprehensive assessment of the contemporaneous and dynamic effects of the ECB's APP across the term structure and their evolution over time. By contrast, other available studies have largely focused on the impact of the initial announcement impact of the APP on asset prices. Second, based on security-level information of asset holdings, aggregate issuance data, and ECB portfolio holdings, we construct a novel granular measure of the “free-float of duration risk,” i.e., the duration-weighted share of public sector debt in the hands of arbitrageurs. Accordingly, our measure tries to mirror the theory set out in Vayanos and Vila (2021). Moreover, we construct projections of that free-float measure, which is a crucial input for the model-based translation of changes in APP purchase parameters into changes in the term premium. To this end, we not only rely on the purchase parameters announced by the ECB, but also account for market expectations by exploiting survey expectations on the path of ECB asset purchases and projecting the market-expected trajectory of reductions in the free-float due to the APP. Third, on the methodological side, we meet the constraints imposed by the relatively short time series of euro area data by deploying a new two-step estimation approach that relies on both fitting the time series of bond yields as well as on utilizing event-based information in the run-up to the ECB's APP. Given this non-standard approach, we also provide a bootstrap procedure to gauge the impact of parameter uncertainty on our estimates.

While our approach utilizes a reduced-form term structure model incorporating the no-arbitrage condition and a stylized version of the duration extraction channel formulated in Vayanos and Vila (2021),

our analysis can inform a more structural modeling of the duration channel of central bank asset purchases. In particular, it could help support the specification and quantification of microfounded equilibrium models.<sup>32</sup>

We acknowledge that several issues relating to modeling, measurement, and the scope of analysis remain to be addressed in future work.

On the modeling side, first, whereas Li and Wei (2013) and this paper serve as examples of how intricate it is to capture the stock effect of QE by itself, there is a case for developing more encompassing models that allow to also incorporate signaling and flow effects in a unified framework. Second, our linear model does not incorporate a lower bound on interest rates. The proximity of rates to the lower bound is potentially relevant for the impact of QE, as documented in King (2019). However, incorporating the lower bound into a term structure model with a duration channel is challenging, especially if there is a need to account for a negative and time-varying level of the lower bound, as would be warranted in the euro area context. Third, another dimension for further refinement is distinguishing the effect QE has on real term premia versus inflation risk premia. While our model captures the impact of QE on the overall nominal term premium, the transmission of QE via the duration channel is likely to affect bond yields mainly via real term premia; see also Kaminska and Zinna (2020) and the references therein. Fourth, it would also be useful—but equally difficult in our specific model framework—to allow for relaxing the assumption of zero risk compensation for supply uncertainty and to build a model that can account for time variation in supply uncertainty in order to capture taper-tantrum type bouts of higher uncertainty. Fifth, focusing in more detail on the implementation of purchases, enhanced models may take into account the impact of purchases on both bond risk premia and liquidity premia.<sup>33</sup>

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<sup>32</sup>For example, King (2015) examines features that are necessary—in particular, with regard to the properties of the implied stochastic discount factor—for general equilibrium models to exhibit a duration channel of the kind we analyze in this paper.

<sup>33</sup>See, e.g., the calibrated search-theoretic model by Ferdinandusse, Freier, and Ristiniemi (2020) as regards the impact of central bank purchases on market liquidity and the relevance of preferred-habitat investors for this channel.

With regard to the measurement duration risk, further work to examine the robustness of the supply variable is warranted. This includes both the analysis of the portfolio rebalancing behavior of different investors and their appropriate assignment to the group of arbitrageurs, and the normalization of the supply variable. With regard to the latter, it would be useful to examine ways to account more directly for the “risk-bearing” capacity of arbitrageurs, for which some measure of investors’ size or capital are worth exploring as alternatives to the total bond supply or GDP.

As for the scope of analysis, our approach could be taken further by also studying the impact of the APP on a wider range of asset classes—in particular, corporate bonds—or by taking a more disaggregated view on the impact across individual euro area countries.<sup>34</sup> Finally, the ECB (and other central banks) have expanded their asset purchases in response to the economic downturn induced by the COVID-19 pandemic. Studying the effect of purchases during that period in an environment of higher risk aversion may grant further insights into how they are transmitted to the yield curve.

## **Appendix. Comparing the Empirical Model to Vayanos and Vila (2021)**

The reduced-form empirical model used in this paper and the U.S. counterpart by Li and Wei (2013) are inspired by the equilibrium model by Vayanos and Vila (2021).<sup>35</sup> As a key feature common to both models, the outstanding supply of debt that has to be borne by arbitrageurs affects the market price of risk of factor exposure and hence term premia. The commonality is best seen when focusing on the model version in Vayanos and Vila (2021), Section 3.2, that switches off time variation in the demand of preferred-habitat investors. Bonds are in zero net supply, and maturity-specific preferred-habitat investor demand is constant, so that  $x^{(\tau)}$  has to be held by arbitrageurs. In the one-factor version of

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<sup>34</sup>See recent work by Costain, Nuño, and Thomas (2022), who present a calibrated model incorporating both a duration extraction and a sovereign credit risk channel to quantify the impact of the ECB’s asset purchases during the pandemic phase on individual sovereign yields.

<sup>35</sup>See also Greenwood and Vayanos (2014).

their model, with short-rate exposure as the only risk, the (instantaneous) expected excess return of holding the  $\tau$ -period bond at time  $t$  is given by  $\mathcal{A}_i(\tau)\lambda_{i,t}$ , where  $\mathcal{A}_i(\tau)$  is the loading of the log bond price on the short-rate factor.<sup>36</sup> The market price of short-rate risk is given by

$$\lambda_{i,t} = a\sigma_i^2 \int_0^T x_t^{(\tau)} \mathcal{A}_i(\tau) d\tau, \quad (27)$$

where  $a$  is the risk-aversion parameter,  $\sigma_i^2$  is the innovation volatility to the short-rate process,  $x_t^{(\tau)}$  is outstanding bond supply of maturity  $\tau$  (i.e., whatever is left net of preferred-habitat holdings) to be absorbed by the arbitrageur, and  $T$  is the maximum maturity of outstanding debt.

In the empirical model used here, the short-term rate is driven by two factors (level and slope). The market price of level/slope risk is given in (20). The time-varying contribution of the supply factor to the factor price is  $\Lambda_{1,ZQ} \cdot Q_t$ , where  $Z$  denotes either level or slope and  $Q$  is our supply variable. This expression is of the same form as in (27), i.e., it is a product of a constant coefficient ( $a\sigma_i^2$  versus  $\Lambda_{1,ZQ}$ ) and a quantity variable ( $\int_0^T x_t^{(\tau)} \mathcal{A}_i(\tau) d\tau$  versus  $Q_t$ ).

Regarding the time-constant multiplier, our reduced-form parameter  $\Lambda_{1,ZQ}$  may hence be interpreted as reflecting risk aversion. At the same time, though, it has to be noted that such a mapping from a structural to a reduced-form model (with more factors) is necessarily incomplete.

Regarding the supply variable, the expression  $\int_0^T x_t^{(\tau)} \mathcal{A}_i(\tau) d\tau$  in Equation (27) can be interpreted as aggregate duration risk in the market.  $\mathcal{A}_i(\tau)$  is the individual bond's (with maturity  $\tau$ ) exposure to short-rate risk. This is weighted by the outstanding bond supply  $x_t^{(\tau)}$  for that maturity and summed up (integrated) across maturities. Greenwood and Vayanos (2014) compare that measure to “simple dollar duration” defined as  $\int_0^T x_t^{(\tau)} \tau d\tau$ , i.e., the weighting is not the bond-specific sensitivity but simply the maturity of the respective bond. Note that the latter expression is analogous to that appearing in the enumerator of our free-float measure  $Q$  in (1), i.e.,

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<sup>36</sup>We slightly adapt their notation to avoid overlaps with the symbols used in this paper.

multiplying maturities with corresponding supply volumes prevailing in those maturity brackets. For Greenwood and Vayanos (2014), this measure of simple dollar duration turns out to be closely correlated to their model-implied (using their parameter calibration) counterparts of short-rate and supply-duration risk.

Summing up, the overall economic mechanism through which quantitative easing affects the term premium is the same in both the equilibrium model of Vayanos and Vila (2021) and the non-structural empirical model used here: an increase in future expected central bank purchases would reduce (current and) expected aggregate duration risk to be absorbed by the market. This reduces the market price of risk, which leads to lower expected excess returns in the future and hence to a contemporaneous compression of term premia and bond yields across maturities. While the equilibrium model and the empirical models share the relevance of bond quantities as a key property, they differ in other details. In particular, their full model (as opposed to the simplified version explained above) features several demand factors determining variation in the demand of preferred-habitat investors. This modeling choice renders the bond volume to be held by the arbitrageur a multi-factor process and these demand factors are also priced in equilibrium. Adopting a factor structure for bond quantities and allowing for pricing of such factor risk is certainly high on the agenda for developing our empirical model further.

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