Does the Cost of Private Debt Respond to Monetary Policy?

Heteroskedasticity-Based Identification in a Model with Regimes

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Abstract

We investigate the effects of the Federal Reserve's quantitative easing and maturity extension programs on the yields of US dollar-denominated corporate bonds using a multiple-regime heteroskedasticity-based VAR identification approach. Impulse response functions suggest that a traditional, rate-based expansionary policy may lead to an increase in yields while quantitative easing is linked to a general and persistent decrease in yields, particularly for long-term bonds. The responses generated by the maturity extension program are significant and of larger magnitude. A decomposition shows that the unconventional programs reduce the cost of private debt primarily through a reduction in risk premia that cannot be entirely accounted for by a reduction in corporate default risk.

Keywords: unconventional monetary policy; transmission channels; heteroskedasticity; vector autoregressions; identification; corporate bond yields.

JEL code: G12, C32, G14

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1. Introduction

The Great Financial Crisis (GFC) has posed unprecedented challenges to the US Federal Reserve (henceforth, the Fed), challenging the conventional wisdom that generally guides monetary authorities (see, e.g., Joyce et al., 2012). Already in December 2008, the zero lower bound for the Fed funds rate was reached, and this forced the Fed to tap its arsenal of alternative, "unconventional" policies. Starting in late 2008, the Fed purchased massive amounts of medium- and long-term assets, primarily agency bonds and mortgage-backed securities (MBS). These policies, which imply an increase of the monetary base through the direct purchase of fixed income securities, are generally known as Quantitative Easing (henceforth, QE). In addition to these outright purchases, in September 2011, the Federal Open Market Committee (FOMC) launched the Maturity Extension Program (MEP). This policy consists of the purchase of long-term Treasury securities financed by the contemporaneous sale of an identical amount of short-term notes previously held in the Fed's balance sheet, for a total worth of $400 billion. The MEP, also known as Operation "Twist" (OT), induced a change in the relative supply of long- and short-term Treasury securities while holding the monetary base constant.

The minutes of the FOMC meeting of December 15-16, 2008 stated that the purpose of these unconventional policies was to "(...) support overall market functioning, financial intermediation and economic growth" since the purchases were expected to "(...) reduce borrowing costs for a range of private borrowers", i.e., the yields of the debt instruments that are issued by firms to satisfy their financing needs, such as corporate bonds. Considering the Fed's objectives and the general importance of debt financing for US firms, it is particularly relevant to understand the effects of conventional and unconventional monetary policies on corporate bonds. To this purpose, our paper uses standard tools from structural vector autoregression (henceforth, SVAR) analysis applying a non-recursive, heteroskedasticity-based identification scheme. In contrast to standard Cholesky schemes, which may lie on thin logical grounds (see, e.g., Gertler and Karadi, 2015), heteroskedasticity-based identification allows the estimation of different on-impact responses in each volatility regime. In essence, we ask whether the idea that a shock to the level and the slope of the Treasury yield curve, such as the one that the monetary authority was able to produce through unconventional policies (as documented by Gagnon et al., 2011; Hamilton and Wu, 2012; Krishnamurthy and Vissing-Jorgensen, 2012) can be effectively transmitted to corporate yields and spreads (at least in a crisis regime) can be empirically supported. As noted by Gilchrist and Zakrajšek (2013), this task is complicated by the simultaneity of policy decisions and movements in the prices of risky financial assets, as well as by the fact that both the corporate yields and spreads (risk premia) targeted by unconventional monetary policy reacted to other common shocks during the financial crisis.

There is an extensive literature that has investigated the ex-post effects (and effectiveness, often performing "bang-for-the-buck" calculations) of QE and the MEP on the securities that were purchased by the Fed and occasionally also on asset classes that were not purchased directly. The majority of the studies found that LSAPs and MEP were successful at lowering the yields on Treasuries as well as on agency debt and mortgage-backed securities, while the effects on assets that were not purchased directly remain more controversial (see, e.g., Justiniano et al., 2012; Rogers et al., 2014; Stroebel and Taylor, 2012).[1] Most of these papers perform event studies supplemented by regressions to control for covariates.

On the contrary, our goal is to investigate the effects of conventional monetary policy, QE, and the MEP on the yields of US-dollar denominated corporate bonds, traded in the US, with different maturities and ratings, in different regimes. As anticipated, on this margin, the literature is thinner and even when evidence has been reported of a significant effect of QE and MEP on the cost of debt to firms, such evidence is weak, with relatively modest medium-term policy multipliers after parameter uncertainty is taken into account, as in Guidolin et al. (2017). Our approach differs from earlier studies in two ways. First, instead of performing event studies, we estimate a regime-switching SVAR model and base our analysis on an identification scheme that relies on heteroskedasticity (IH) proposed by Rigobon (2003) and Lanne and Lütkepohl (2008). This choice is crucial as pointed out by Herwartz and Plodt (2016), who used simulations to show that IRFs identified using (co)variance shifts offer the most precise measures of the actual dynamics.[2] Second, differently from the bulk of the literature, we pursue an a-priori investigation, instead of an ex-post assessment (for instance, see in Gagnon et al., 2011) of the effects of monetary policies on securities that were not directly included in the purchase programs but which appear to be the target of policy-makers, simulating the effects of each type of policy through an impulse response function (IRF) analysis.

Even in the absence of a formal model, our results are compatible with the idea that *unconventional* policies lead to the desirable responses (i.e., a reduction in the cost of capital for firms, including those of lower credit standing). These effects are precisely estimated in all the regimes under analysis, in contrast to the weaker evidence reported in earlier work (e.g., by

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[1] Krishnamurthy and Vissing- Jorgensen (2012) and Gagnon et al. (2011) found that LSAP1 induced a decline in the 10-year Treasury yield of 100 bps and 91 bps, respectively. LSAP2 was studied, among others, by Greenwood and Vayanos (2014), Krishnamurthy and Vissing-Jorgensen (2012), Meaning and Zhu (2011), and Swanson (2011) who reported a reduction in Treasury yields in the range 15-55 bps. Hamilton and Wu (2012) have documented that MEP reduced Treasury yields with a maturity in excess of 2 years by about 17 bps, while short-term yields increased by a similar amount.

[2] The performance of the identification via IH depends on the relative size of the volatility shifts (that should be large for accurate identification) and the length of the sample (to exceed 200 observations). Our application fulfills the requirements in Herwartz and Plodt to support an IH identification scheme.

Guidolin et al., 2017) that use traditional, recursive identification schemes. However, also in our analysis, the estimated effects turn out to be stronger in a regime that is identified as describing a state of crisis and market turmoil. In our third, more volatile regime, the effect of a QE-type shock (an *unanticipated* decline in the 10-year rate) is a statistically significant decline of up to 14 bps and 45 bps for investment grade and non-investment long-term yields, respectively. In the case of MEP shocks (an *unanticipated* decline in the 10-year rate accompanied by an *unanticipated* increase in the 1-year T-bill rate), the corresponding responses are a decrease of about 22 and 107 basis points for investment and non-investment grade corporate yields, respectively. Further analysis within an SVAR identified through a similar heteroskedasticity scheme shows that most of these hefty effects come from a precisely estimated reduction of credit spreads over time, i.e., from a risk premium channel, resorting to the nomenclature proposed by Longstaff (2010). Additional tests based on the monthly series of Gilchrist and Zakrajek's (2012) Expected Bond Premium (henceforth, EBP) show that such effects cannot be entirely imputed to a reduction in corporate default risk.

Our analysis shows that a *conventional*, expansionary monetary policy would have led to a generalized increase in corporate yields; such a perverse reaction of the cost of private debt might have been sizeable, e.g., an increase by about 80 bps for non-investment grade and of the order of 20 bps for investment grade bonds, in the crisis state. This is in line with the findings of Guidolin et al. (2017) and can be interpreted as a consequence of the inflationary expectations that rate-based policies may trigger (see Ang and Piazzesi, 2003). Alternatively, the increase in corporate yields may be triggered by the policies bringing the rates below Brunnermeier and Koby's (2018) "reversal rate", i.e., the rate at which accommodative monetary policy reverses its intended effects and becomes contractionary, as a result of banks' asset re-valuation from duration mismatch being more than offset by decreases in net interest income.

These results hold after performing a throughout set of robustness checks, which involve introducing common shocks in the SVAR model, and replacing the series of corporate yields with credit spreads. Indeed, IRFs estimated from an SVAR applied to Treasury term spreads and corporate credit spreads leads to similar empirical estimates as those obtained from the baseline model, in contrast to earlier findings that especially MEP policies would affect corporate yields but would cause mixed effects on risk premia (see, e.g., Guidolin et al., 2017).

Three related papers are Gilchrist and Zakrajšek (2013), Guidolin et al. (2017), and Kontonikas et al. (2020). Gilchrist and Zakrajšek have estimated the effects of the Fed's QE and MEP programs on corporate credit risk by also employing a heteroskedasticity-based, event study approach. While they report that QE announcements led to a significant reduction in CDS spreads for both investment- and speculative-grade corporate bonds, their paper focuses on estimating the impact of policy announcements and does not produce a full IRF analysis. Guidolin et al. (2017) simulate the effects of a range of monetary policies in flexible VAR models with regimes, identified using a range of Cholesky decompositions. They report that the responses of corporate bonds to unconventional monetary policies are statistically significant and of the sign intended by policy-makers only when implemented in the regime identified with the crisis state. However, despite the considerable robustness checks performed, whether or not the ordering implied by Cholesky identification schemes remains contentious. We depart from their analysis by using an alternative identification scheme and by performing a detailed set of robustness checks on the effects of common shocks to the interest rate series trigged by changes in general economic conditions.[3] Kontonikas et al. (2020) analyze the effects of monetary policy shocks on contemporaneous corporate bond returns and use a variance decomposition framework to disentangle the role of interest rate vs. discount rate. Although the connection between corporate bond returns and yields is easy to grasp so that there is a clear relationship between our work and Kontonikas et al.'s, their analysis emphasizes the role played by Fed fund rate shocks only and therefore rules out the focus on unconventional policies that characterizes our research.

The rest of the paper is structured as follows. Section 2 introduces the data and describes the questions under investigation. Section 3 reviews the methodology. Section 4 reports our key empirical findings. Section 5 performs a set of robustness checks while Section 6 concludes.

2. The data

The main issue with an analysis of bond market prices is that these markets are generally far less liquid and transparent than, for instance, equity markets (see Bessembinder and Maxwell, 2008). A large portion of the literature relies on indices of corporate bond yields for research purposes (see, for example, Neal et al., 2000; Longstaff, 2010). One of the drawbacks of this practice is that the resulting indices include both callable and non-callable bonds, according to time-varying weights that reflect their market values. The callability feature (i.e., the option for the issuer to refund a bond before the stated maturity) may have a significant impact on the behavior of the bond yield, we manually construct corporate bond portfolio yield series, relying on transactions reported and collected by the Trade Reporting and Compliance Engine (TRACE), a system managed by the Financial Industry Regulatory Authority (FINRA). We merge the information on US-dollar denominated corporate bond trades with the details concerning the bond issues (their

[3] The existence of regimes in the (long-run) relationships between interest rates (official, credit and government debt), monetary aggregates (money stock and monetary base) and real income has been recently exploited by Olmo and Sanso-Navarro (2015) to test the existence of a shift in the transmission mechanism through which the unconventional policies affected the economy. However, their focus is not explicitly on the effectiveness of different types of unconventional monetary policies on yields.

rating) retrieved from the Mergent Fixed Income Securities Database.[4]

The final step of the filtering process consists in classifying the observations according to maturity and rating: a bond is considered to be short-term (ST) if, at the time of the recorded trade, its remaining time to maturity is less than five years and long-term (LT) otherwise. Concerning the rating clusters, bonds are classified as investment grade (IG) if their rating is higher or equal than A- (in Standard & Poor's and Fitch scales) or A3 (Moody's scale) and non-investment grade (NG) otherwise. Next, we construct four series of weekly yields, one for each possible combination of the assigned cluster for rating and maturity, to obtain the following series: investment grade short-term bonds (IGST), investment grade long-term bonds (IGLT), non-investment grade short-term bonds (NGST), and non-investment grade long-term bonds (NGLT). The weekly yield series for each portfolio are constructed by averaging the yields for all the bonds traded in each week belonging to that portfolio.[5] The final results are four weekly (Friday-to-Friday) yield series for a sample October 1, 2004 - March 30, 2017.

The riskless US yield curve is summarized by four weekly series of constant maturity Treasury yields: 1-month, 1-, 5-, and 10-year yields. Those series are retrieved from the Federal Reserve Economic Data (FRED of Saint Louis) for the same sample period as above.

3 Methodology

3.1 Baseline model and standard identification schemes

We assume the model has the following structural form

$$\boldsymbol{A}\boldsymbol{y}_t = \boldsymbol{c}_0 + \boldsymbol{B}_0\boldsymbol{y}_{t-1} + \boldsymbol{\varepsilon}_t, \tag{1}$$

where $\boldsymbol{y}_t$ is an 8x1 vector collecting the 1-month Treasury yield ($1\text{m}T_t$), the 1-year yield ($1\text{y}T_t$), the 5-year yield ($5\text{y}T_t$), the 10-year yield ($10\text{y}T_t$), and the IGST, IGLT, NGST, and NGLT yields defined as above.[6] Details are available upon request. $\boldsymbol{B}_0$ captures the lagged effects of the endogenous variables y_t.[7] The 8x8 matrix **A** has diagonal elements equal to one, while its off-

[4] We consider U.S. Corporate Debentures, Corporate Medium Term Notes, and U.S. Corporate MTN Zeros.

[5] For some trades, the yield is missing in the TRACE repository: in that case we use the coupon rate, payment frequency, issue date, and remaining time to maturity to calculate the yield using standard formulas. We have also tried to build portfolios weighted by their outstanding amounts and found qualitatively similar results in terms of their means and standard deviations of the portfolio yields.

[6] Results are insensitive to replacing $1\text{m}T_t$ with the effective Federal funds rate.

[7] As it is customary, a VAR(p) with p > 1 can also be represented in companion form as a VAR(1) by simply expanding the vector of endogenous variables to include lags of $\boldsymbol{y}_t$ up to p -1. In the following, we work with a VAR(1) representation while being aware that a higher-order VAR may be easily accommodated. Although a 8-variable system is by no means small, higher-dimensional systems have been considered, e.g. in Bernanke et al. (2005), that contain more variables than just the interest rates that we entertain in this paper. Resorting to a ninth common factor variable in Section 5.1 takes steps in this direction. There is also a literature that has extended such large systems to study the effects of the sacrifice ratio between growth

diagonal elements capture the contemporaneous interactions across endogenous variables. The vector $\boldsymbol{\varepsilon}_t$ contains the structural form white noise shocks, which are assumed to have zero mean and variance $\sigma^2_{\varepsilon,i}$ $\forall i$ = {1, ..., 8} and to be orthogonal, both contemporaneously and across time, i.e., $E(\varepsilon_{i,t}, \varepsilon_{j,t}) = 0, \forall i \neq j,\ E(\varepsilon_{i,t}, \varepsilon_{j,t'}) = 0\ \ \forall i \neq j, t \neq t'$.
The reduced form associated with the structural model is

$$\boldsymbol{y}_t = \boldsymbol{A}^{-1}\boldsymbol{c}_0 + \boldsymbol{A}^{-1}\boldsymbol{B}_0\boldsymbol{y}_{t-1} + \boldsymbol{A}^{-1}\boldsymbol{\varepsilon}_t = \boldsymbol{c}_1 + \boldsymbol{B}_1\boldsymbol{y}_{t-1} + \boldsymbol{\eta}_t, \tag{2}$$

in which the residuals are related to the structural residuals by $\boldsymbol{\eta}_t = \boldsymbol{A}^{-1}\boldsymbol{\varepsilon}_t$. It is well-known that the starting point for the identification of the structural parameters is to estimate the reduced form model in (1) by OLS, obtaining estimates of $\boldsymbol{c}_1$, $\boldsymbol{B}_1$, and the covariance matrix of the reduced form residuals (call it $\boldsymbol{\Omega}_\eta$); then, the structural coefficients might be retrieved from the reduced form estimates. Of course, if **A** were known, this would be sufficient to recover the structural coefficients. In our case, because the covariance matrix of the reduced form residuals has 36 elements, we have only 36 equations for 64 unknowns (i.e., the 8 diagonal elements of the covariance matrix of the structural form residuals, given our assumption of zero correlation across structural shocks, plus the 56 off-diagonal elements of the matrix **A**). Hence, there are more unknowns than equations, which means that a continuum of solutions exists.

Various methods of identification have been used in the literature: sign or exclusion restrictions on some parameters, which ideally should derive from economic theory, have been often employed, but such restrictions generally remain untestable. A commonly used method is Cholesky's decomposition, which imposes that the matrix **A** is triangular, which in our application implies zero-restrictions on the contemporaneous effects among variables to achieve exact identification. Unfortunately, the typical Cholesky restrictions, when applied to a model that contains asset prices, tend to be implausible and, as shown by Ehrmann et al. (2011), a standard Cholesky identification may fail to achieve the proper identification because of the strong asymmetries that it forces on the VAR system. As a result, in this paper we adopt an identification scheme based on the heteroscedasticity (in short, IH) that tends to characterize financial data.

3.2 Identification through heteroskedasticity

Rigobon (2003) provides the theoretical derivation of the IH methodology and its application in the form of a GMM estimation methodology. The methodology is useful in any situation in which it is difficult to impose credible exclusion restrictions, like in the one we describe in the following. The precise form of heteroskedasticity is not crucial in that framework. The key idea is that

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and inflation. For instance, Gross and Simmler (2019) use regime switching SVARs to contrast the effects of conventional and unconventional monetary policies.

breaks/regimes in the reduced-form error covariance matrix can be associated with changes in the on-impact responses of the variables to the shocks, which in turn reflect in instabilities in the identified impulse response functions (IRFs) across volatility regimes. In our case, we shall assume an SVAR model with $S > 1$ regimes in the covariance matrix of the structural form.[8] Such regimes help because they imply that each additional heteroskedastic regime adds more equations than unknowns in the system. Because the lack-of identification can be pinned down in the need for 28 additional restrictions, but each regime provides 36 of them, clearly $S = 2$ is sufficient to exactly identify our VAR model.

For instance, when there are two regimes in the variances of the structural shocks, the sample can be split into two subsamples, presenting high and low volatility, respectively. Assuming that the structural parameters are stable across the two regimes, we obtain

$$\boldsymbol{A}\boldsymbol{y}_t^s = \boldsymbol{c}_0 + \boldsymbol{B}_0\boldsymbol{y}_{t-1}^s + \boldsymbol{\varepsilon}_t^s \qquad s = 1{,}2, \tag{3}$$

where $Var[\boldsymbol{\varepsilon}_t^s] = \boldsymbol{\Sigma}_\varepsilon^s$, and $\boldsymbol{\Sigma}_\varepsilon^1 \neq \boldsymbol{\Sigma}_\varepsilon^2$ with the covariance matrices of the reduced form shocks given by $\boldsymbol{\Omega}_\eta^1 \neq \boldsymbol{\Omega}_\eta^2$**.** In this new system of equations, the unknowns are 72, i.e., the $(N^2 - N) = 56$ elements of **A**, the $N = 8$ elements of $\boldsymbol{\Sigma}_\varepsilon^1$, and the $N = 8$ elements of $\boldsymbol{\Sigma}_\varepsilon^2$. Given the assumptions above, we can estimate a reduced form covariance matrix for each subsample: each covariance matrix provides $N + (N^2-N)/2 = 36$ elements. With two regimes, they are exactly 72 in total, solving the identification problem.[9]

We implement a recursive definition of regimes based on the residuals of the estimated reduced-form VAR model and computing time-varying, rolling-window variances over 12-week samples for each variable: we identify a shift in regime every time for one or several endogenous variables their relative variance (of the VAR residual series) exceeds its historical sample average plus one standard deviation by at least a third of the standard deviation for a minimum of 24 consecutive weekly observations. Using this procedure, we identify three separate regimes, both in the baseline cases and when exogenous shocks are considered.

In our application, the system is identified by the heteroskedasticity existing in the data, but still this is only true up to a rotation of the matrix **A.**[10] Since all such rotations of **A** allow us to

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[8] As shown in Rigobon (2003), the estimates of the conditional mean coefficients are consistent, regardless of how the heteroskedasticity is modelled, provided the number of regimes has been correctly specified.

[9] When the data exhibit S heteroscedasticity regimes, the system that has $S[N + (N^2-N)/2]$ equations (one covariance matrix per regime) and $N^2 - N + SN$ unknowns (N structural variances for each regime, plus the parameters of **A**). A solution is guaranteed for $S[N + (N^2-N)/2] \geq N^2 - N + SN$, which is satisfied for $S \geq 2$. In particular, the system is exactly identified in presence of two regimes. Otherwise, the system has more equations than unknowns and the additional equations provide testable over-identifying restrictions.

[10] A rotation is the multiplication of the matrix **A** by another matrix that is full rank and has determinant equal to one. Since both the matrix **A** and its rotation solve the system, the problem of how to differentiate the two is solved in practice by imposing additional exclusion or sign restrictions to force upon the methodology the selection of a unique rotation.

solve the system of equations, the sign restrictions we discuss below have the objective to ensure that we pick an economically meaningful rotation (see Herwartz and Plodt, 2016, for details). Specifically, the sign restrictions considered in our baseline model are as follows: since the Treasury term structure is upward sloping (a feature that will be confirmed on average also by our data, see Section 5, but also see the robustness check in Section 5.3), we assume that an increase in the yield of short term Treasuries has a positive effect on all other bonds with longer maturities, both Treasuries and corporate, which is generally consistent with the expectations hypothesis of the yield curve. Formally, given that in our case $\boldsymbol{y}_t \equiv [1mT_t \;\; 1yT_t \; 5yT_t \; 10yT_t \; IGST_t \; IGLT_t \; NGST_t \; NGLT_t]'$ and

$$\boldsymbol{A} \equiv \begin{bmatrix} 1 & \tau_{12} & \tau_{13} & \tau_{14} & \alpha_{15} & \alpha_{16} & \alpha_{17} & \alpha_{18} \\ \tau_{21} & 1 & \tau_{23} & \tau_{24} & \alpha_{25} & \alpha_{26} & \alpha_{27} & \alpha_{28} \\ \tau_{31} & \tau_{32} & 1 & \tau_{34} & \alpha_{35} & \alpha_{36} & \alpha_{37} & \alpha_{38} \\ \tau_{41} & \tau_{42} & \tau_{43} & 1 & \alpha_{45} & \alpha_{46} & \alpha_{47} & \alpha_{48} \\ \alpha_{51} & \alpha_{52} & \alpha_{53} & \alpha_{54} & 1 & \rho_{56} & \rho_{57} & \rho_{58} \\ \alpha_{61} & \alpha_{62} & \alpha_{63} & \alpha_{64} & \rho_{65} & 1 & \rho_{67} & \rho_{68} \\ \alpha_{71} & \alpha_{72} & \alpha_{73} & \alpha_{74} & \rho_{75} & \rho_{76} & 1 & \rho_{78} \\ \alpha_{81} & \alpha_{82} & \alpha_{83} & \alpha_{84} & \rho_{85} & \rho_{86} & \rho_{87} & 1 \end{bmatrix} \tag{4}$$

so that the τ parameters indicate the contemporaneous spillovers across Treasury yields of different maturities, the ρ parameters the spillovers across corporate yields in different rating and maturity clusters, and the α parameters the spillovers across Treasury and corporate yields. Although we have experimented with alternative sets of restrictions (results are available upon request), our baseline case consists of imposing $\tau_{21} < 0$, $\tau_{31} < 0$, $\tau_{41} < 0$, $\tau_{32} < 0$, $\tau_{42} < 0$, and $\tau_{43} < 0$.[11] Because we believe that this causality sequence should apply both to the direct effects of shocks on short-term rates (as measured by the matrix **A**) and the overall effects, including indirect spillovers (as measured by $\mathbf{A}^{-1}$), we impose the equivalent set of restrictions on $\mathbf{A}^{-1}$.[12]

3.2.1 Common shocks and estimation methodology

There is one additional, necessary condition to achieve identification, besides the existence of heteroskedasticity, i.e., that in spite of the time-varying variances, the structural shocks remain uncorrelated. This assumption, which is crucial as it implies that each additional heteroskedastic regime adds more equations than unknowns in the system, may not be fulfilled if those residuals are driven by one or more common shocks.[13] Such common shocks are very likely when asset

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[11] The sign of the restrictions are negative but this is not inconsistent with standard economic meaning because the matrix **A** pre-multiplies the endogenous variables on the left-hand side of (1).

[12] The sign restrictions limit the space in which parameters have to be searched to minimize the moment restrictions. This influences the speed of convergence but does not affect precision unless the estimates are on the boundaries. As we will see later, very few of the coefficients are on the boundaries.

[13] Bacchiocchi and Fanelli (2015) show that identification may be achieved even allowing the VAR matrices to change across regimes as well. However, this requires imposing additional exclusion restrictions on **A**,

prices (here, interest rates) are considered because markets are tightly interconnected, even contemporaneously. The explicit introduction of common shocks allows us to model the possible covariance among residuals and thus to safely assume the orthogonality of the structural residuals after adequate transformations. In presence of common shocks, $\boldsymbol{z}_t$, our model has the following structural and reduced forms:

$$\boldsymbol{A}y_t = \boldsymbol{c}_0 + \boldsymbol{B}_0\boldsymbol{y}_{t-1} + \boldsymbol{D}_0\boldsymbol{z}_t + \boldsymbol{\varepsilon}_t \tag{5}$$

$$\boldsymbol{y}_t = \boldsymbol{A}^{-1}\boldsymbol{c}_0 + \boldsymbol{A}^{-1}\boldsymbol{B}_0\boldsymbol{y}_{t-1} + \boldsymbol{A}^{-1}\boldsymbol{D}_0\boldsymbol{z}_t + \boldsymbol{A}^{-1}\boldsymbol{\varepsilon}_t = \boldsymbol{c}_1 + \boldsymbol{B}_1\boldsymbol{y}_{t-1} + \boldsymbol{D}_1\boldsymbol{z}_t + \boldsymbol{\eta}_t, \tag{6}$$

where $\boldsymbol{D}_0$ captures the effect of the exogenous common shock vector. It may also be plausible to consider the lagged effects of the common shocks, $\boldsymbol{z}_{t-1}$. We conduct a robustness check concerning the assumption of orthogonality of the structural form residuals in Section 6.1, introducing a common shock $\boldsymbol{z}_t$, both in its contemporaneous and lagged specification.

All the common shocks are assumed to have zero correlation among them and with the structural shocks. The variances of shocks are $\sigma^s_{z,k}$ and $\sigma^s_{\varepsilon,n}$, respectively. In our case, we have N = 8 equations and K = 1 common shock, the number of equations is given by the covariance matrix in each regime, i.e., $S[N + (N^2 - N)/2]$ in total, while the unknowns are the elements of matrix $\mathbf{A}$, i.e., $(N^2 - N)$, the elements of matrix $\mathbf{D}_0$, i.e., $K(N - 1)$, the variances of the common shocks in each state, i.e., KS, and the variances of the structural shocks in each regime, i.e., NS. In general, (see Rigobon, 2003), the system is identified if and only if, the number of states satisfied $S[N + (N^2 - N)/2] \geq N^2 - N + K(N - 1) + SK + SN$ or

$$S \geq 2\frac{(N+K)(N-1)}{N^2 - N - 2K} \tag{7}$$

and if one additional regime in the covariance matrix adds more equations than unknowns, i.e., the number of common shocks satisfies $N^2 - N - 2K > 0$, then $K < (N^2 - N)/2$. In our case with $N = 8$ and $K = 1$, equation (7) is satisfied. This means that with the introduction of one or two common shocks (as lagged common shocks), identification is exactly achieved in the presence of three heteroskedastic regimes, but the ability to perform over-identifying tests disappears.

For the purpose of our analysis, it is reasonable to consider that the structural residuals of our model might be influenced by the prevailing general business conditions. For that reason, we employ the Aruoba-Diebold-Scotti (ADS) business conditions index to measure common shocks.[14] The ADS business conditions index is aimed at tracking real business conditions: the average value of the ADS index is zero and it assumes larger (positive) values to indicate better-

which may be undesirable in our application, or—which is our case—that the number of regimes exceed the strict minimum to achieve identification, here $S = 2$.

[14] The ADS index is based on a number of observable economic indicators: weekly initial jobless claims, monthly payroll employment, industrial production, personal income less transfer payments, manufacturing and trade sales, and quarterly real GDP, see Aruoba et al. (2009).

than-average conditions or lower (negative) values to signal worse-than-average conditions. Thus, the ADS index is suitable to obtain a measure of time-varying business cycle conditions. The ADS index is updated continuously as new data are released, which is at least once a week: we use the vintages available for our sample period (from October 1, 2004 to March 30, 2017) from the website of Federal Reserve Bank of Philadelphia.

Once the number of regimes has been identified, we estimate the parameters of interest by minimizing the following minimum distance function

$$\min_{\boldsymbol{\theta}\in\mathbb{C}} \sum_{s=1}^{3}\left\{\boldsymbol{A}'\boldsymbol{\Sigma}_{\varepsilon}^{s}\boldsymbol{A}-\boldsymbol{\Omega}_{\eta}^{s}\right\}'\left\{\boldsymbol{A}'\boldsymbol{\Sigma}_{\varepsilon}^{s}\boldsymbol{A}-\boldsymbol{\Omega}_{\eta}^{s}\right\} \tag{8}$$

where $\mathbb{C}$ is the sub-space of values for $\boldsymbol{\theta}$ such that the sign constraints specified under equation (5) are satisfied, $\boldsymbol{\theta}$ collects the parameters to be estimated, $\boldsymbol{\theta}\equiv vec(\boldsymbol{A})'$, $\boldsymbol{\Sigma}_{\varepsilon}^{s}$ is the regime-specific diagonal matrix of variances of the structural shocks, and $\boldsymbol{\Omega}_{\eta}^{s}$ is the covariance matrix of the reduced-form residuals estimated in each regime s = 1, 2, 3. The criterion function is replaced by $\left\{\boldsymbol{A}'\boldsymbol{\Sigma}_{\varepsilon}^{s}\boldsymbol{A}-\boldsymbol{D}_0{}'\boldsymbol{\Sigma}_{z}^{s}\boldsymbol{D}_0-\boldsymbol{\Omega}_{\eta}^{s}\right\}'\left\{\boldsymbol{A}'\boldsymbol{\Sigma}_{s}\boldsymbol{A}-\boldsymbol{D}_0{}'\boldsymbol{\Sigma}_{z}^{s}\boldsymbol{D}_0-\boldsymbol{\Omega}_{\eta}^{s}\right\}$ when there are common shocks, where $\boldsymbol{\Sigma}_{z}^{s}$ is the covariance matrix of the common shocks in regime s, which is assumed to be a diagonal matrix with elements $\sigma_{z,k}^{s}$. This is analogous to a GMM estimator (see Rigobon, 2003), the distribution of which is easily derived, at least asymptotically. Interestingly, the consistency of estimates holds even when the data contain *more* regimes than the ones specified, while consistency when the data contain less regimes may be more problematic.

To avoid reporting inferences that solely rely on asymptotics, we block-bootstrap the *p*-values of our parameter estimates. For each of the heteroskedasticity regimes, we use the estimated regime-specific covariance matrices to create new data with the same covariance structure in each bootstrap replication. For each draw, we estimate the coefficients by minimizing the moments given the restrictions. We use 5,000 bootstrap replications throughout.

3.3 Impulse response policy experiments

We follow the tradition in applied macroeconomics and use impulse response functions (henceforth, IRFs) to quantitatively track between h= 1 and h = 52 weeks (since the inception of one or more shocks) and understand the effects of monetary shock of different types on corporate bond yields. However, when identification is supported by regimes that occur only in the covariance matrix of the structural residuals, a unique matrix of the contemporaneous, structural effects (along with the dynamic VAR matrices that are assumed to be constant over time) implies time-homogeneous IRFs (see Bacchiocchi and Fanelli, 2015).

It is possible to extend the concept of IRF to non-linear models by defining a generalized IRF (see, e.g., Koop et al., 1996), with reference to a regime-switching framework that also involves matrices of parameters that enter the conditional mean function of the SVAR. In our case, we assume that the conditional mean dynamic parameters ($\mathbf{B}_1$ and $\mathbf{D}_1$, where present) are in no way dependent on the variance regimes that simply enter in the identification of the SVAR. However, the residual covariance matrices $\boldsymbol{\Omega}_{\eta}^{s}$ s = 1, 2, 3 depend on the regime and offer the opportunity to solve three distinct local problems, in which also the matrices of the contemporaneous cross-variable effects have been made regime-dependent.

An h-step-ahead IRF is then defined as

$$IR_{\Delta\boldsymbol{\varepsilon}}(h) = E[\boldsymbol{y}_{t+h}|\boldsymbol{y}_t(\boldsymbol{\varepsilon}')] - E[\boldsymbol{y}_{t+h}|\boldsymbol{y}_t(\boldsymbol{\varepsilon})]\,, \tag{9}$$

where the sample path $\boldsymbol{y}_t(\boldsymbol{\varepsilon}')$ is equal to the sample path $\boldsymbol{y}_t(\boldsymbol{\varepsilon})$ with the exception of the initial value of $\boldsymbol{y}_t$, which has been perturbed by a shock $\Delta\boldsymbol{\varepsilon}$ (see Potter, 2000). This definition can be generalized to fit a regime-switching framework as:

$$IR_{\Delta\boldsymbol{\varepsilon}}(h,s) = E\left[\boldsymbol{y}_{t+h}\middle|S_{t,}\boldsymbol{\varepsilon}_t + \Delta\boldsymbol{\varepsilon}; \boldsymbol{y}_{t-1}\right] - E\left[\boldsymbol{y}_{t+h}\middle|S_{t,}\boldsymbol{\varepsilon}_t; \boldsymbol{y}_{t-1}\right]. \tag{10}$$

In a framework with regimes, an h-step-ahead IRF depends on the state S_t prevailing at time t when the shock occurs because it affects the estimated matrix of contemporaneous effects, $\boldsymbol{A}_s$. Notably, because we analyze models with regimes in which the VAR matrix is not regime-dependent, we only need information about the state at the time of the shock. Even though more complex approaches are possible (Monte Carlo techniques to simulate the ergodic distribution of regimes, or assuming equal probabilities across regimes), in this paper, we assume that the regime prevailing when the shock occurs is known, which is consistent with Sections 3.1 and 3.2. Because IRFs are computed using estimated coefficients, they clearly also reflect estimation error. Accordingly, we construct confidence intervals for the IRFs using bootstrapping techniques (see Killian, 1998). An impulse response function is considered statistically significant if zero is not included in the bootstrapped confidence interval.[15]

To simulate the three monetary policies we are interested in, we consider their effects on Treasury rates (as suggested by Bernanke and Blinder, 1992). This approach relies on two main assumptions: (i) in the conventional monetary policy case, Fed funds rate and 1-month Treasury yield tend to strongly co-move; (ii) in unconventional monetary policies case, QE, and MEP, ought to be at least able to lower Treasury yields. We make these choices for two reasons: first, the traditional approach of measuring conventional monetary expansion by changes in monetary aggregates fails to recognize that the rate of growth of monetary aggregates also depends on the trend of growth of the currency component of the money supply (see Bernanke and Mihov, 1998); second, in the case of the MEP, the policy does not imply the creation of monetary base, but,

[15] Additional details concerning the bootstrapping exercise can be found in a working paper version of the manuscript at https://papers.ssrn.com/sol3/papers.cfm?abstract_id=3458310.

instead, it induces a change in the relative supply of long- and short-term Treasuries so that it would be impractical to simulate these policies relying on shocks to monetary aggregates.

Therefore the premise of this paper is that both conventional and unconventional policy shocks can be approximated as changes in nominal Treasury yields. While in the case of Fed fund rate-driven policies this represents a persuasive choice, in the case of QE and the MEP, Treasury yields are just intermediate targets of the monetary policy, and one may wonder whether it is realistic to assume that a one-time unconventional policy shock may be causing a large effect on Treasury rates. There is a growing literature trying to pin down the quantitative impact of Fed actions on the Treasury yields with rather sparse findings. The estimated effects vary between approximately 5 bps (as in Gagnon et al., 2011) and 70 bps (as in D'Amico and King, 2013) per week per 100 billion of QE intervention, and between 3 and 15 bps per 100 billion of MEP intervention (see Hamilton and Wu, 2010; Swanson, 2011). Given that the average size of LSAP 1–3 may be quantified in approximately US450$ billion, this corresponds to a shock in the range 22-315 bps (assuming the policy had not been anticipated). Similarly, because the MEP consisted of a plan to purchase US$400 billion of Treasury securities with remaining maturities of six through 30 years and to sell an equal amount of Treasuries with maturities of three years or less, this maps, assuming the policy change had been unexpected, into a shock between 12 and 60 bps.

In practice, using the 1-month bill yield to represent short-term rates and the 10-year Treasury yield to represent long-term rates, we simulate a conventional monetary expansion through a negative shock to the 1-month T-bill yield. Furthermore, we represent QE through a negative shock to the 10-year Treasury yield; finally, we simulate the MEP through a negative shock to the 10-year Treasury yield accompanied by a simultaneous positive shock to the 1-month yield. As in most of the literature, the shock considered is equal to one standard deviation.[16] However, the within-regime linear nature of the model makes adjustments to account for the actual size of the policy shocks feasible. With reference to a similar set of experiments and by looking at the published estimates in the literature, Guidolin et al. (2017) report averages of the weekly impact on 10-year Treasury yields per a standardized 100-billion USD policy intervention that range between 10 and 200 bps for LSAP1 and between 25 and 300 bps for LSAP2 and settle on a 100 bps impact effect in their bootstrap exercise. In the case of the MEP, they estimate impact effects on long-term yields, standardized to a 100 billion policy, that fall between a few bps only and as much as 30 bps.[17]

[16] The IRFs are not altered by adding (simultaneous and lagged) common shocks, as the latter are not perturbed by the policy impulse(s) and thus they cancel out.

[17] Such shock sizes compare to an estimated one-standard deviation of errors equal to 15 bps in the case of 10-year Treasuries and of 4 bps in the case of 1-month T-bills in the crisis regime. The estimated error standard deviations are substantially lower in regimes 1 and 2. This means that the baseline results in

4 Empirical results

4.1 Regime definition

We use the reduced form residuals from an estimated VAR(1) to define the heteroskedastic regimes, as described in Section 3.1, and define three regimes.[18] Because the consistency of the GMM estimator holds when the number of regimes in the data is under-identified vs. the true but unknown data generating process, a parsimonious selection of S =3 seems appropriate.[19] Figure 1 reports evidence on the nature of the three heteroskedasticity regimes. Regime 1 is the regime recurring more often: it accounts for about 65% of the observations in the sample. In this regime, the estimated variances of the structural shocks (see Table 1) are lower than in the other regimes (with the exception of 10-year Treasury yield shocks, which display a lower variance in regime 2). Although the regime-specific means of the yields are in no way used in their definition, we also note that the means of the residuals of both corporate bonds and Treasuries in this regime are lower than in the overall sample. Thus, regime 1 can be considered a frequently occurring and tranquil state characterized by low expected yields on all bonds.

Regime 2 is the least frequent regime: it accounts for less than 6% of the observations in our sample and appears to have occurred in the 2006-2007 period (see Figure 1). In this regime, residual variances are structurally higher than in regime 1 (excluding the cases of NIGLT bonds and 10-year Treasuries) but lower than in regime 3. Also the means are higher in this regime vs. regime 1 for all yield series, whereas the standard deviations of the series are *lower*. The simultaneous occurrence of higher residual variances with lower variances of the series indicates that the R-squares implied by our VAR(1) filter structurally decline in this regime. Thus, we consider regime 2 as a transient state, essentially preceding the crisis state, in which bond markets are characterized by high yields, low volatilities, and widespread unpredictability.

Regime 3 is characterized by the highest variances of residuals (see Table 1), except for IGLT bonds. In particular, the variance of the residuals of the series of NIG yields skyrockets. This regime accounts for about 30% of the observations in our sample, including the credit crunch

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Section 4 may grossly under-estimate the potential quantitative effects of QE. Separate calculations that asymmetrically increase 1-month T-bills by one standard deviation but decrease 10-year Treasury rates by twice their standard deviation to represent a MEP-style shock also confirm that Section 4 provides a substantial under-estimation of the potential effects derived in our exercise, especially in the crisis regime.

[18] To select the proper order of the VAR model, we use the Bayesian Information Criterion (BIC). When p ranges between 0 and 10, we select a VAR(1) model. The VAR(1) turns is stable and hence covariance stationary. The OLS estimates of the model are reported in an Appendix available upon request.

[19] We formally test the occurrence of the regimes in the reduced-form variances through a standard LR Chow-type test. We focus on the null hypothesis that the VAR reduced-form residual variances are constant across the three regimes, $\mathbf{\Sigma}_{\varepsilon}^{1} = \mathbf{\Sigma}_{\varepsilon}^{2} = \mathbf{\Sigma}_{\varepsilon}^{3}$ for a total of 72 restrictions. The null hypothesis of constant covariance parameters is starkly rejected.

phase during the 2008 crisis and 2010-2011, when short-term Treasuries reached the zero lower bound and the warning of a US debt ceiling-driven crisis dominated the news. The periods of the LSAPs in March 2009, the MEP, and LSAP3 are all captured by this regime. For what concerns the means and standard deviations in Tables 3 and 4, the average corporate yields are higher than in the other two regimes (excluding the IGST series, which has a higher mean yield in regime 2), while the means of yields on Treasuries are lower vs. regime 2. Regime-specific standard deviations are the highest in this regime for corporate bonds (especially for NIG paper), while they are similar to regime 1 as far as Treasuries are concerned. Thus, regime 3 is a crisis state driven by the fact that during market crashes, investors tend to move towards safer assets (i.e., Treasuries and, to some extent, IGST bonds), lowering their yields, while causing the yields on high-risk bonds (i.e., non-investment grade corporates) to soar.

4.2 Estimation and identification through heteroskedasticity

The estimated matrix **A** is reported in Table 2: on the basis of the bootstrapped p-values for the t-test applied to the individual coefficients reported in parenthesis, the entire matrix is highly statistically significant. The coefficients measuring contemporaneous effects among corporate bonds are substantial: the positive spillovers from corporate bond yields of shorter maturities to the longer-term ones in the same rating cluster are the strongest effects we report.

The highest spillover effects involving Treasuries are instead those from the shortest maturities to longer-term bonds, which is an implicit feature of the upward sloping term structure of Treasuries under the Expectations Hypothesis. Spillovers across Treasury and corporate bond yields are generally weaker, even though all of them are precisely estimated. Exceptions are represented by the negative contemporaneous effect among short-term bonds, in particular the 1-month T-bill rates and the yields of IGST. This empirical result appears to be consistent with the preferred-habitat theory since bonds in the same maturity segment should be considered substitutes in terms of investment choices to a specific preferred-habitat investor.[20]

Under S = 3 regimes, Section 3 shows that a heteroskedasticity-based identification strategy leads, when N = 8 and p = 1, to 28 over-identifying restrictions. The model fails to be rejected from a formal (quasi) LR test: under the null hypothesis consisting of the 28 restrictions,

[20] The preferred-habitat theory has been recently formalized by Vayanos and Vila (2009). While the expectation hypothesis states that the presence of risk-neutral agents implies that the term structure is determined only by current and expected short rates, the preferred-habitat theory states that markets are segmented and the relative supply of assets influences their yields for each specific maturity. In this framework, preferred-habitat investors have a strong preference for a specific maturity segment and demand only bonds that correspond to their maturity habitat.

the LR statistic is equal to 35.04 and yields a p-value of 0.169 (under a $\chi^2(28)$ distribution).

As a last step, we test whether the dynamics of the VAR are also time-varying through a test of H_0: $\boldsymbol{A}_1 = \boldsymbol{A}_2 = \boldsymbol{A}_3$ but $\boldsymbol{\Sigma}_\varepsilon^1 \neq \boldsymbol{\Sigma}_\varepsilon^2 \neq \boldsymbol{\Sigma}_\varepsilon^3$ vs. the alternative that $\boldsymbol{A}_1 \neq \boldsymbol{A}_2 \neq \boldsymbol{A}_3$ and $\boldsymbol{\Sigma}_\varepsilon^1 \neq \boldsymbol{\Sigma}_\varepsilon^2 \neq \boldsymbol{\Sigma}_\varepsilon^3$ (implicitly, the regimes defined by heteroskedasticity must be aligned with those for the conditional mean function). In both cases, the sign restrictions that follow equation (5) and that uniquely define a rotation of the $\boldsymbol{A}_s$ matrices have been imposed. In this case, the (quasi-)LR test becomes equals 144.49 and allows us to reject the null with a p−value of 0.021 (under a $\chi^2(112)$ distribution). We thus find formal support that the three breaks involve both the conditional mean and conditional covariance matrix of the VAR model.

4.3 *Effects of a conventional monetary expansion*

We now turn to the computation of the IRFs for the three types of shocks representing a conventional monetary expansion, the QE program, and the MEP, respectively. Figures 2-4 show the IRFs of corporate yields of all rating and maturity clusters in each of the three regimes and over a response interval of 52 weeks along with bootstrapped 95% confidence intervals.

Starting with the analysis of the implications of a conventional expansionary monetary policy, simulated as a negative one standard deviation shock to the 1-month Treasury yield (approximately equal to 41 bps). In contrast with the expected effects of an expansionary monetary policy and with its stated objectives by policy-makers, our analysis shows a generalized *increase*, rather than a decrease, in corporate rates.[21] In particular, the effect is stronger for NIG bonds, it leads to a maximum impact within the first month from the implementation of the policy, and the result appears to hold not only in the crisis state (where it is the strongest), but mildly also in the tranquil and transient regimes. In terms of persistence, any statistically significant effects declines rapidly, reaching zero after four-to-six months.

In the crisis regime 3, this perverse effect is stronger in terms of magnitude in all cases, similarly to Dahlhaus's (2017) findings; the most significant increases in yields are those recorded for short-term bonds: in the periods following the implementation of the policy, NIG bonds are subject to an increase in a range of 60-80 bps, whereas the maximum impact on IG yields is around 22 bps, recorded in the second week after the shock occurs. As far as long-term yields are concerned, the estimated effects are significant only in the fifth week after the policy shock in the case of IG bonds (a positive increase of 13 bps is registered), while for NIG yields, a peak increase of 74 bps appears after 4 weeks.

[21] Even in normal times, the Fed might be using non-borrowed reserves in addition to short rate impulses to implement conventional policies. Moreover, in spite of their high correlations, 1-month rates are an imperfect proxy for shocks to the Fed funds rate. We abstract from these aspects in our analysis.

In both regimes 1 and 2, a conventional, expansionary policy has positive and significant effects on IGST yields, inducing an increase in rates of the order of 1-5 bps. IG long-term bonds are not characterized by a significant increase in yields. NIG yields in both maturity clusters are subject to a statistically significant increase in regime 1 and 2 of approximately 4 and 15 bps, respectively.

Our findings are consistent with Guidolin et al. (2017), who found that the effects of conventional monetary policy on corporate yields may carry the wrong sign, given the desiderata of policy-makers. These results should be taken with caution, given-their policy implications in terms of the signals to the fixed income market about future inflation and growth implied by an accommodative policy. In fact, in times of crisis, a Fed funds rate reduction that is transmitted to the rate of 1-month T-bills, is likely to signal to the market a concern by the Fed about future growth and employment, and this may generate negative expectations about the future demand. Coherently with the findings by Bernanke and Mihov (1998), our results are consistent with a response of corporate yields to an expansionary shock that reflects the economic regime prevailing at the time when the shock occurs. Kontonikas et al. (2020) have reported that, during the 2007-2009 financial crisis, which is mostly captured by our regime 3, the elevated uncertainty led to an increase in riskier (i.e., non-investment grade) yields, while policy rates were being sharply cut, which is a result in line with the positive and significant IRF estimated for NIG bonds.

A few theoretical papers have recently argued that regimes may exist in which *accommodative*, rate-based policies may have a *contractionary* impact on corporate bond yields. Gourio (2013) argues that a non-linear effect of policy rate cuts on corporate yields may also be a consequence of the inflationary expectations that this policy may trigger through a classical monetarist channel. Because conventional policies increase the monetary base, these are expected to be transmitted over time to prices and hence increase nominal interest rates; the expectation of a monetary tightening creates business cycle risks that explain the increase in corporate rates. Eggertsson, Juelsrud, and Wold (2017) have found that negative nominal interest rates may not be truly expansionary owing to zero-interest-bearing cash becoming relatively more competitive. Moreover, we conjecture that because of the structural impairment in US credit markets caused by the GFC, Brunnermeier and Koby's (2018) "reversal interest rate"—the rate at which accommodative monetary policy reverses its intended effects and becomes contractionary in terms of lending costs and supply—may have been crossed in 2008-2010, the periods of our sample characterized by regime 3. Brunnermeier and Koby discuss how this effect may occur when banks' asset re-valuation from duration mismatch is more than offset by decreases in net interest income on new business (see, e.g., Covas, Rezende, and Vojtech, 2015; Egly, John, and Mollick, 2018), lowering banks' net worth and tightening their capital constraints, which are the conditions likely to be engulfing the US credit markets between 2008 and 2010.[22]

4.4 *Effects of quantitative easing*

We have analyzed the effects on corporate yields of the QE program, simulated as a one standard deviation negative shock to the 10-year Treasury yield, which is approximately 61 bps on impact; this estimate falls at the lower end of the 22-315 bps range that has been reported in the literature with reference to the LSAP1-3 programs. The responses to QE are shown in Figure 3 and are remarkably different across maturities. Long-term corporate yields significantly decline, while short-term yields record a significant increase: in particular, the increase occurs after about six months from the implementation of the policy in the case of IG bonds and in the period immediately after the shock in the case of NIG bonds. In general, responses to QE-type policies are more persistent vs. the ones obtained for the case of a conventional monetary expansion. For that reason, we also estimate the IRFs over a horizon exceeding the 52-weeks reported in Figure 3 and find that any effects decline in their magnitude and remain statistically significant for approximately 70 weeks after a QE shock.

The effect on IGLT yields is a statistically significant decrease in all regimes considered: 14 bps in the crisis state and about 6 bps in both regimes 1 and 2. The same holds for NIG bonds of the same maturity (excluding the first period). In that case, the maximum decline in the crisis state is 45 bps, whereas in the tranquil regime, the yield declines by 20 bps, and in the transitory regime by 16 bps. All these effects are precisely estimated and economically sizeable. The response of NIGST yields in all regimes is an initial significant increase (in the order of 10 bps), followed by a persistent decrease, lasting about 6 months, up to 8 bps, 6 bps, and 18 bps in the tranquil, transitory, and crisis states, respectively. In the case of IGST corporates, QE seems to produce insignificant effects in all regimes during the initial 10 months. Subsequently, yields increase in all regimes (especially in the crisis state). Nevertheless, the magnitude of this effect slowly fades after period 52 and stops being precisely estimated.

Excluding the case of IGST bonds, the effects we report are consistent with the policy objective of reducing the borrowing costs to firms. Notably, the strongest effects of QE are

[22] The finding that the undesirable effects of rate cuts would be weaker and imprecisely estimated in correspondence to regime 1 (that characterizes the sub-period 2011-2016) is fully consistent with the finding in Brunnermeier and Koby (2018) that over time, after the end of the GFC and the normalization of the functioning of the US credit markets, the reversal interest rate would slowly increase as asset revaluation fades out and fixed-income holdings mature: most of regime 1 may be characterized by rates close to zero but yet large enough to fall close to this endogenous lower bound without causing contractionary effects, also because the net interest margins and the overall profitability of US banks did improve starting in 2012-2013 (see Tran, Lin, and Nguyen, 2016). Ampudia and Heuvel (2018) show that banks' profitability response to interest rate cuts is non-monotonic – in normal times, interest rate cuts increase banks' valuations, although this does not hold in low-rate environments.

obtained in a regime of crisis, which is precisely when the policy is most likely to be implemented/needed, even though the objective of reducing the cost of debt faced on average by firms could also be achieved (probably, at the cost of having to scale up the size of the QE operations) also in regimes 1 and 2.

Compared to the IRF estimates in Guidolin et al. (2017), our effects are larger in magnitude for IGLT bonds and smaller for NIG bonds. This difference might be caused by the simpler—possibly problematic—identification strategy applied by Guidolin et al., i.e., a Cholesky triangularization in which the variable ordering plays a crucial role and IG bonds are causally prior to NIG ones, so that it is exactly IG corporate papers that appear to be the most influenced by shocks to other variables. Although this may be consistent with a preferred habitat hypothesis for interest rates, the lack of robustness of their conclusions to the identification strategy should alert policy-makers of the fact that QE strategies may be effective exactly at times where these are most needed, i.e., in times of crisis, when the goal is to lower the cost of capital of firms of less-than perfectly credit quality.[23] More generally, the literature that has investigated the effects that QE has produced on risky yields of assets different from those purchased directly by the Fed (e.g., Krishnamurthy and Vissing-Jorgensen, 2011; D'Amico and King, 2013) and reported that QE was effective in reducing their yields during periods of turmoil. However, our analysis also reveals that the effects of QE are statistically significant and carry the desirable sign in non-crisis regimes, indicating that this policy should steadily enter (as it seems to have, see Farmer and Zabczyk, 2016, and recent reactions to the pandemics by the Fed) the toolkit available to policy-makers, regardless of the GFC-related events that did lead to promote these policies in late 2008.

4.5 Effects of the maturity extension program (Operation "Twist")

Figure 4 reports the IRFs concerning corporate yields when the shock mimics the effects of the MEP, which we simulate as a one-standard-deviation negative shock to the 10-year Treasury yield accompanied by a positive one-standard-deviation shock to the 1-month Treasury yield. In general, the effects generated by MEP are precisely estimated and carry the desired sign as it triggers a reduction in yields in every regime, in particular in the crisis one. However, the yield responses turn less persistent than in the case of QE in Figure 3. For instance, in the case of the IGST yield, the response lasts a shorter period, and the rate decline is of about 2 bps, 4 bps, and 22 bps in the tranquil, transitory, and crisis regimes, respectively. Declines of a similar magnitude are estimated for IG long-term yields, although their responses are more persistent. The effects of the MEP on NIG yields are significant and larger than those of any other policy: short-term NIG

[23] Guidolin et al. (2017) perform robustness checks on the Cholesky ordering and report generalized IRFs, but the tendency of the response of high-grade to exceed that of low-grade rates turns out to be pervasive.

yields decline by up to 70 bps in the crisis state, 9 bps in the tranquil regime, and 15 bps in transitory regime; in the case of long-term yields, significant effects appear after a few periods from the occurrence of the shock and the overall decline reaches about 22 bps in the tranquil regime, 28 bps in the transitory state, and 107 bps in the crisis state. These are hefty policy effects.

Differently from the case of QE in Figure 3, the IRFs to MPE shocks show the expected signs and are significant in a rather homogeneous fashion. This can be explained by the peculiar nature of the MEP, which is a policy aiming at reducing the slope of the Treasury yield curve without increasing the monetary base (see Hamilton and Wu, 2010), thus without inducing expectations of an increase in inflation. The results in Figure 4 contribute to shed further light on the findings in Figure 3 concerning the QE program, as QE does imply an expansion of the monetary base and a likely impact on inflation expectations (see, e.g., Bernanke and Mihov, 1998). Moreover, in regime 3, the results appear to be qualitatively in line with those in Gilchrist et al. (2015) and Guidolin et al. (2017) but of stronger magnitude, which is of course of practical relevance and in accordance to Mishkin's (2009) conjecture that monetary policy may be more effective just in times of crisis. However, the IRFs in Figure 4 also reveal non-negligible and precisely estimated effects in regimes 1 and 2, which is not the case in Guidolin et al.'s work, where an arguably weaker (because arbitrary) identification scheme had been applied.

Our finding that that the MEP may be more effective than QE is partly surprising, although earlier papers had assessed the MEP to be a remarkably effective strategy (e.g., see Hamilton and Wu, 2010) through a duration risk channel, to be contrasted to a simpler monetary base one. However, we should recall that our empirical assessment is only focussed on the cost of private debt and assumes effects on Treasury rates that are data-driven. It would be tempting to jump to the conclusion that in the case of QE, the resulting expansion of the monetary base may have acted to significantly offset the effectiveness of QE so that the FED might have achieved better results sterilizing its purchases. Nonetheless, we need to remember that the feasible size of MEP finds a natural size limit in the initial amount of bills and short-term notes in the FED's portfolio, which means that MEP is a useful but not all-purpose weapon in a central bank's arsenal.

5 Discussion and robustness checks

5.1 The effects of common shocks

The introduction of common shocks in the model allows us to relax the assumption of orthogonality of the structural residuals, as explained in Section 4. We start by introducing one common shock and estimating the VAR(1) model in (7) when the common shock concerns the ADS business conditions index. By applying the same methodology as in Section 5 (also as far the selection of p = 1), we find that the regime definition based on the dynamic, rolling window residual variances leads to the same number and characterization of the regimes as in Section 5 and Table 1. In fact, a comparison of Tables 3 and 1 reveals negligible differences, of 10% at most in the largest estimated residual variances.

The estimated covariance stationary VAR(1) is tabulated in an Appendix. The results turn out to be similar to those in the baseline model in terms of significance, sign, and magnitude of the estimated conditional mean coefficients. Interestingly, more than a half of the coefficients loading on the ADS index growth rate are significant; in general, a high growth rate in the business cycle index forecasts higher Treasury yields (the only exception concerns the 10-year Treasury, for which the estimated coefficient fails to be significant) but lower corporate yields. The direction of these effects can be explained thinking that a strong ADS index indicates above-average economic conditions, which may induce investors to move from safe investments (such as Treasuries, whose price declines and yields climb up) to investments in riskier assets (such as corporate bonds, especially of NIG type) which therefore display lower, subsequent rates. In this perspective, the weaker effects on 10-year Treasury rates are sensible, as long-term government bonds imply high duration risk.

The GMM estimates of **A** obtained from the identification-through-heteroskedasticity methodology for this model specification is shown in Table 4, panel A. Interestingly, also the estimated matrix of contemporaneous effects is hardly affected by the introduction of common shocks so that Table 4 resembles Table 2. The key exception involves NIGST yields, which are now positively influenced by Treasury yields of all maturities. Another substantial difference is that shorter-term Treasuries (1-month and 1-year) are negatively influenced by the yields of NIGST bonds and IG long-term bonds, respectively, albeit the magnitude of the effect is relatively small (and yet precisely estimated).

Given the estimated parameters presented above, we compute the IRFs for the three policies studied in Section 5 and report the results in Figure 5. The effects produced by a conventional monetary expansion are the same as before for yields of all rating-maturity clusters and regimes, in terms of the sign, magnitude, significance, and persistence: in regimes 1 and 2, the effects are modest and usually not statistically significant; in regime 3, the responses turn perverse (although estimates tend to be imprecise, also because short-term rates were hardly actually used to affect corporate rates in the periods that are best characterized as regime 3) in the sense that an expansionary policy would push corporate bond yields higher and hence discourage investments and production. In the case of the IRFs that refer to the QE programs, the responses have the expected signs and tend to be precisely estimated in regime 3, especially for long-term corporate bonds, both NIG and IG, and over a period of 5-6 months. The effects are weaker in regimes 1 and 2, although often precisely estimated, which would give policy-makers reasons to pursue QE policies also in moderately volatile economic environments. Finally, and consistently with the results in Section 5, the MEP leads to the strongest yield responses, with the same qualitative patterns as QE but with considerably higher estimates, across all regimes. When MEP is concerned, also IGST corporate yields are affected, but the effect tends to disappear in 2-3 months after the shock. However, all these remarks closely mimic those reported in Section 5.

We repeat the GMM estimation and the IRF analysis including a second common shock that represents the lagged effect of the growth rate in the ADS index. Apart from minor details, the empirical definition of the regimes for identification purposes and the variances of structural residuals are identical to those presented in Tables 3 and 4 (panel A). An Appendix reports detailed estimation results for the case of both contemporaneous and lagged common shocks and shows that results are quite homogeneous in two cases apart from the fact that, when lagged common shocks are taken into account, the coefficients loading the ADS index on bond yields are reduced in magnitude and significance compared to the specification without a lagged common shocks. The estimated matrix **A** is shown in panel B of Table 4: coefficients are almost identical to those estimated in the model with no lagged common shocks, and the precision of the estimates remains high. An Appendix available upon request computes the IRFs for the three policies studied in this paper from the model with contemporaneous and lagged common shocks. The IRFs are practically unchanged vs. Figure 5 and require no additional comments.

In conclusion, we obtain evidence that the introduction of one single common shock is sufficient to ensure that the assumption of orthogonality of the structural residuals is satisfied. Nevertheless, the matrices **A** and $\mathbf{B}_0$ obtained under this specification are similar to those retrieved from the baseline model, so that the responses estimated by the IRFs are analogous. However, this alternative specification allows us to isolate the embedded inflationary expectations effects of QE. Indeed, the IRFs now describe declining responses—especially for IGST yields—to shocks, in contrast to MEP, that instead fails to trigger strong inflationary expectations reactions and hence turns out to be more effective over medium-long term horizons.

5.2 The effects of identification through heteroskedasticity

It is natural to ask whether identification through heteroskedasticity leads to empirical results that differ in important ways vs. those obtained under traditional identification schemes, based on imposing an ordering on the variables. Interestingly, given the limitations of identification through ordering/triangularization—especially in the face of the jolt represented by the GFC that has altered the variance of shocks—we cannot really speak of the benefits of an identification scheme based on second moments without resorting to artificial simulations (see Herwartz and Plodt, 2016). However, in Figure 6, we perform a visual comparison of the state-dependent IRFs derived from our identification strategy (as in Figure 3) with those derived using the same variable ordering (within a standard Cholesky scheme) as in Guidolin et al. (2017).[24] Three remarks are in order. First, a simple look at the left axis scales reveals that in the crisis regime, all estimated effects are massively larger (often double), although of the same sign, under a heteroskedasticity-based identification scheme. Without attempting to flesh out a theoretical justification for those stronger effects, it is intuitive that a more effective identification ought to lead to sharper results in terms of their size and the precision of the resulting estimates. Second, the only type of corporate paper for which, again in the crisis regime, there is a visible difference in the shape of the IRF as a function of the horizon, is NIGST bonds: the effects are U-shaped and precisely estimated only for horizons in excess of 3 months in our paper, but monotonically increasing and significant at horizons up to 5 months under a more traditional identification scheme. Third, although these are all flat and hard to tell apart, all the IRFs for the pre- and post-crisis regimes (in Figure 6, we adopt the same nomenclature for regimes as in Guidolin et al.'s) seem to be approximately the same under the two schemes. All in all, adopting the identification strategy supported in our paper seems to produce a first-order quantitative effect exactly in the regime in which the unconventional monetary policies had been introduced.

5.3 Excluding Periods of Non-Upward Sloping Yield Curves

Because the sign restrictions in (5) assume an upward sloping yield curve, the corresponding rotation may be problematic in a portion of our sample. Therefore, we proceed to classify each weekly observation in our 2004-2017 sample on the basis of whether the corresponding riskless yield curve (as defined by the 1-month, 1-, 5-, and 10-year Treasury maturities) were monotonically upward sloping or not. We can identify an uninterrupted period (January 2006 – September 2007) in which the risk-free term structured failed to be upward sloping, and it was instead frequently downward sloping. In fact, we observe that most of these 21 months of data are characterized by the prevalence of what we have identified as regime 1. Therefore, excluding these 90 observations is almost equivalent to restricting our analysis to two regimes only. Using what is essentially a two-regime model, we have re-estimated the model for the Treasury and corporate yields and computed the resulting IRFs.

Figures 10 and 11 show that, as far as regimes 2 and 3 are concerned (we do not report IRFs for regime 1 because these would be based on less than 30 observations and estimated very

[24] Guidolin et al.'s triangular factorization follows two criteria: variables to be shocked, i.e. long, medium and short Treasury rates, are placed on top of the ordering; the rest of the variables are ordered on the basis of their residual maturity, with Treasuries preceding corporate bonds. Figure 6 concerns the effects of a one standard deviation shock to 10-year Treasury yields that want to mimic the effects of QE, but a similar figure has been produced in the case of the MEP and is available upon request from the authors.

imprecisely), our key results are intact.[25] In Figure 10, QE shocks produce quantitatively similar and precisely estimated impacts on corporate yields, especially long-term ones, as in Figure 3. In Figure 11, the MEP shocks lead to quantitatively similar and precisely estimated impacts on corporate yields in the case of the crisis regime 3 as those found in Figure 4.

5.4 Spread analysis

To gain additional insights on how different types of monetary policies may affect corporate bond yields and hence affect real investment decisions, we have also analyzed the response of bond yield spreads: policy measures may affect rates in two ways, either by changing the risk premia required to buy bonds or by altering the value of time factor, i.e., the riskless interest rate. Hence, a policy may be effective on the latter (and we know that both QE and to some extent also MEP did reduce the riskless interest rates at the long-end of the maturity spectrum, while conventional policies do that by construction at the short-end) but cause perverse effects on bond risk premia. Alternatively, QE and MEP could have been successful in stimulating the economy not only by lowering the general level of interest rates but also by inducing a reduction in the compensation demanded by investors for expected default risk and, more generally, in the average price of bearing exposure to corporate credit risk. This increase in investors' risk appetite—by lowering the price of default risk—may have placed additional downward pressure on corporate borrowing rates and further stimulated business fixed investment.

For each recorded corporate bond trade recorded in TRACE, the credit spread is calculated as the difference between the yield of the corporate bond and the yield of a Treasury bill or note (that we use as a proxy for the risk-free rate) with a maturity which approximately matches the remaining life of the bond.[26] Next, we average the spreads of all the bonds traded each week for each of the four portfolios defined in Section 3. This methodology based on the trade-by-trade calculation of spreads is more accurate than simply taking the difference between corporate yield averages and Treasury rates (see Lin and Curtillet, 2007, for a discussion).

An Appendix available upon request shows the estimated parameters from a (covariance stationary) VAR(1) model for corporate and Treasury spreads:[27] more than half of the

[25] A set of plots available upon request shows that especially in the crisis regime 3 and with reference to short-term bonds, expansionary conventional rate-based policies keep leading to undesirable effects on the cost of debt of corporations, even though most of the response functions fail to be precisely estimated. We thank one anonymous referee for suggesting this robustness check.
[26] We discretize the time-to-maturity of each corporate bond in the following categories, by selecting the closest from the list: 1 month, 3 months, 6 months, 1 year, 2 years, 3 years, 5 years, 7 years, and 10 years. Then, we match such imputed maturity with Treasuries of similar maturity traded in the market.
[27] The model contains eight endogenous variables, i.e., the 1-month T-bill rate along with seven corporate and government bond yield spreads. Therefore 4 of the spreads concern corporate bonds and are obtained

coefficients are significant. As in the baseline model for the bond yields, the corporate spread series display a significant forecasting power for each other, while the NIG corporate spreads have significant predictive power for Treasuries. The resulting regime definition appears to be robust to replacing the yield with the spread series: regime 1, i.e., the tranquil state, is characterized by low mean rates but intermediate average spreads; regime 2, i.e., the transient state, is characterized by high mean rates and low mean spreads; regime 3, i.e., the crisis state, is associated to medium mean rates but high mean spreads, as typical of a crisis state in which the Fed aggressively cuts short-term rates as a first-order measure. The estimated residual variances are ranked across regimes in the way already found in Section 5, with the variance of the residuals increasing when we move from regime 1 through 3. The matrix of residual variances of the structural innovations is shown in the Appendix 5 and it is essentially identical to Table 2, with reference to the baseline model case. The estimated matrix **A** in Table 5 implies that the major difference compared to the baseline case is that the long-term Treasury spread displays a positive contemporaneous association with the spread of IG bonds, both short- and long-term ones. This means that the entire Aa-Aaa cluster moves together in terms of differentials vs. riskless, matching bonds.

We proceed with the estimation of the IRFs, which are reported in Figures 7-9. In general, the results obtained for corporate credit spreads are in line with those discussed in Figures 2-4. Yet, there are two meaningful exceptions to this general finding. The effect of a conventional monetary expansion on the IGLT spread is larger than the response of the IGLT spread in all three regimes, especially in the crisis state. While the difference may be due to imprecise estimates in the two applications, in this perspective, monetary policy has a large impact on risk premia, which is reassuring. Second, the sign of the responses of the IGST spreads are now more sensible in response to the QE program than they were in total terms, as the peak effect in absolute terms occurs at around the 15th week (a decrease of 8 bps); for longer horizons, the IRF increases and stops being precisely estimated. In summary, the effects on the spreads are similar to those on yields in our baseline model. Results are clearer in terms of the direction of the responses for IGST and, in general, of a larger magnitude. As in the baseline case, a conventional, expansionary monetary policy triggers an *increase* in corporate spreads, which is consistent with policy rates reacting to adverse news to economic fundamentals, which signal a deterioration in the outlook for credit quality and reflect a downward revision to future growth prospects more than an increase in discount rates. As a result, the credit spread may increase while longer-term risk-free

in the way described in the main text and the remaining three are term spread inferred from the term structure of riskless interest rates, computed as a difference between long-term and 1-month yields.

rates decline.[28] In fact, there is an empirical literature that has reported a non-monotone relationship between credit spreads and Treasuries (e.g., Duffee, 1998) and suggests that conventional measures may struggle in reducing the cost of capital of firms.

In any event, there is no doubt that unconventional monetary policy in the US could be expected to affect corporate bond risk premia in desirable directions. If any, there is evidence that such effects are generally and persistently sensible as they go in the appropriate directions, occasionally even more than it is the case for the overall level of yields, probably reflecting more complex effects on other, relevant risk premia that however our data cannot account for.[29]

5.5 Expected bond risk premium analysis

Gilchrist and Zakrajek (2013) have proposed a decomposition of credit spreads in the Expected Bond Premium (henceforth, EBP), i.e., the component of spreads not explained by default risk (proxied by observable characteristics) and the default risk-driven portion. Using their monthly data, we have estimated a version of our baseline model (based on 150 observations for the period Oct. 2004 – March 2017) that also included Gilchrist and Zakrajek's EBP. In Figure 12, we report the estimated IRFs from the baseline three-regime model triggered by conventional and unconventional monetary policy shocks. The regimes are defined as in Figure 2 of the baseline exercise. The plots show that in all regimes, a rate-based conventional monetary policy fails to cause precisely estimated responses by the EBP.[30]

The findings in Figure 12 support those in Figures 8 and 9 concerning QE and MEP shocks: in the crisis regime 3 and, at least to some extent, in the post-crisis regime 2, the EBP is substantially lowered by the unconventional monetary policies. For instance, QE reduces the premium by 20 bps in the initial months and the effect slowly converges to zero, being still at more than 10 bps after 1 year. The effect is smaller compared to Figure 8 (between 50 and 25 bps, declining over time) but precisely estimated, an indication that roughly 40% of the stimulus effect that goes through corporate spreads comes from a pure reduction in risk premia and the remaining 60% from a decline in priced default probability risk. Similarly, a MEP shock reduces the premium by 30 bps in the initial months and the effect slowly converges to zero, being still

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[28] This is consistent with the empirical findings in Kontonikas et al. (2020) for the GFC period and with the result in Figure 6 that the negative relationship between changes in short rates and spreads is most pronounced for lower-rated corporate credits, a segment of the market that was especially vulnerable to adverse macroeconomic shocks during the early stages of the GFC.

[29] During a crisis, IGST bonds may become illiquid and this may increase their yields above the level justified by a decline of the riskless rate as well as of the default risk premia induced by QE.

[30] This is more realistic than the results in Figure 7 on the total, weekly credit spreads on corporate bonds, in which credit spread would be shooting up in a significant matter for several weeks after the initial shock. The wide confidence bands attached to the regime 1 IRF in the plots derive from the fact that in this monthly analysis, regime 1 just characterizes 20 observations out of 150 (a fraction of the sample similar to what is reported in Figure 2, but for weekly data) which leads to imprecise estimates.

estimated at almost 25 bps after 1 year. The effect is much smaller if compared to Figure 9 (between 100 and 35 bps reduction over time) but precisely estimated, an indication that, initially, roughly 30% of the stimulus effect comes from a pure reduction in risk premia; this fraction grows to approximately 2/3 as the horizon lengthens.

6 Conclusions

In this paper, we have investigated to what extent monetary policies may affect the cost of private debt, for instance, as expressed by the yields implied by traded corporate bonds, in times of financial crises and volatile asset prices. This question is key to policy-makers and for the very understanding of the performance of modern market-based economies. We use statistical methods that exploit the very conditions of distress and volatility typical of financial crises to achieve a causal identification of monetary policy shocks following the heteroskedasticity-based identification methodology proposed by Rigobon (2003). In particular, we have tested the effects of conventional and unconventional monetary policies on the yields of US corporate bonds at different tenors and for two key rating classes.

Even though the absence of a fully developed quantitative theoretical model advises caution in extending the Granger-causal connections informing the estimated IRFs to policy advice, the dynamic responses of private debt to the simulated unconventional policies are statistically significant in all regimes. Notably, such responses are always stronger in what we have labeled as the crisis state, i.e., the regime in which the (residual) volatilities are the largest. As shown by Gourio (2013), the pricing of a corporate bond reflects not only the risk linked to the credit-worthiness of the issuer but also a risk connected to a disaster probability. Because, during a crisis, this component may increase, in contrast with the stated objectives of policy-makers, a conventional, expansionary monetary policy may lead to a generalized increase, rather than a decrease, of corporate yields. The responses to QE appear instead to be strongly persistent and go in a desirable direction. The effect on IG long-term yield is a significant decrease of 14 bps in the crisis state. The same holds for NIG yields of the same maturity, for which the decline reaches 45 bps. Excluding the case of short-term private debt, the effects of QE are consistent with the objective of reducing the borrowing costs for firms, and results echo Gilchrist et al.'s (2015). The responses generated by MEP are instead larger, generally significant and in line with the intended direction, as the program triggers a reduction in yields. For IG bonds, the decrease is about 22 bps in the crisis state. The effects of the MEP on NIG yields are larger than QE's: short-term yields decrease up to 70 bps and long-term yields up to 107 bps in regime 3.

Our results are robust to the introduction of common shocks and to replacing the series of total corporate yields with credit yield spreads, which measure ex-ante risk premia, or to a monthly analysis of Gilchrist and Zakrajek's (2012) EBP, the "pure" risk premium portion of the spreads. The IRFs estimated from a model for spreads show similar patterns vs. those obtained from the baseline model, in contrast with the mixed findings in Guidolin et al. (2017): a relatively agnostic identification scheme based on the deep properties of financial data, i.e., time-varying variances, escapes the puzzling implication that monetary policies would be weakly effective because they may lead to contradictory reactions by the risk premia in US corporate bonds.

However, QE and the MEP may also have changed investors' expectations about future Fed rates through a signaling effect. For instance, Krishnamurthy and Vissing-Jorgensen (2011) estimated the signaling effect through the magnitude of shifts of forward rates and showed that this accounts for a significant portion of the decrease in 10-year bond rates deriving from QE. In our analysis, we have not tried to disentangle the pure unanticipated effects of QE and MEP that exploit the segmented nature of the US bond market from the signaling, systematic components of monetary policy, even though this remains an interesting venue for further research.

Our analysis has adopted an approach in which unconventional monetary policies have been measured by their estimated effects on the US riskless yield curve. Although this is backed by abundant empirical evidence, this assumes the existence of an effective transmission channel of QE and MEP to Treasury rates. Finally, in this paper, we have only considered the effects of US monetary structural policy shocks on US corporate yields and spreads but omitted the fact that in the aftermath of the GFC, a wave of expansionary unconventional policies have been enacted by most key central banks worldwide. Recent work by Groba and Serrano (2020) finds a strong connection between foreign monetary policy and the default risk of US and European firms over the pre-crisis period. It would be interesting to extend our work to include foreign policy shocks.

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