JOURNAL OF FINANCIAL AND QUANTITATIVE ANALYSIS Vol. 59, No. 2, Mar. 2024, pp. 626–651

doi:10.1017/S0022109022001557

Monetary Policy and Bond Prices with Drifting Equilibrium Rates

Carlo A. Favero
Bocconi University Department of Finance, Innocenzo Gasparini Institute for Economic Research (IGIER), and Centre for Economic Policy Research (CEPR)
carlo.favero@unibocconi.it (corresponding author)

Alessandro Melone
The Ohio State University Fisher College of Business
amelone.research@gmail.com

Andrea Tamoni
Rutgers Business School
atamoni@business.rutgers.edu

Abstract

We study the drift and cyclical components in U.S. Treasury bonds. We find that bond yields are drifting because they reflect the drift in monetary policy rates. Empirically, modeling the monetary policy drift using demographics and productivity trends, plus long-term inflation expectations, leads to cyclical deviations of bond prices from their drift that predict bond returns in- and out-of-sample. These bond cycles can be interpreted as term premia or/and temporary deviations from rational expectations in a behavioral framework. Through the lens of our model, we detect a significant role of the latter in determining the cyclical properties of yields with short maturities.

I. Introduction

Bond prices are codrifting: They are nonstationary and they share a common trend.[1] Understanding the drift in bond prices is of essential importance in the current scenario in which several fiscal authorities are considering issuing very long-dated bonds or pension funds would like to buy long-term Treasuries for duration matching purposes ("Need Discount Debts? Try 50-Year Bonds," *Wall Street Journal*, Jan. 28, 2021); indeed it is at the long end of the curve where the

We are grateful to Pierluigi Balduzzi, Hendrik Bessembinder (the editor), John Cochrane, Vito Gala, Kasper Joergensen, Scott Joslin (the referee), Fabio Trojani, Christian Wagner, Josef Zechner, and participants at the LTI/Bank of Italy Workshop 2021, the BFWG Conference 2021 at Queen Mary University of London, the IAAE Annual Conference 2021, the 2021 INQUIRE Residential Seminar, and the 2022 Asset Pricing Workshop for their comments. Alessandro Melone acknowledges the financial support from the Austrian Science Fund (FWF).

[1]Bond prices have been drifting in the last 40 years because their secular drivers have been drifting. As we shall see later, we find that the age structure of the population, potential output growth, and long-term inflation expectations jointly capture the stochastic trend in yields. Throughout the article, we use the words trend and driver interchangeably.

implications for a drifting term structure become more relevant. This article proposes a model of monetary policy with drifting equilibrium rate that describes in a coherent way the codrifting term structure. Our empirical analysis shows that, once the common trend has been removed, the cyclical (i.e., stationary) component of yields emerges as a strong predictor of excess bond returns.[2]

The time-series nature of bond prices makes it essential for term structure factor models and (empirical models built on) monetary policy rules to account for the nonstationarity of bond yields, a fact that has been acknowledged (e.g., Kozicki and Tinsley (2001)) and modeled within an arbitrage-free dynamic term structure model (DTSM) with a shifting endpoint (e.g., Bauer and Rudebusch (2020)). Despite these contributions, the exact nature of the drivers of the stochastic trend in yields, the relation between the cyclical components of yields with the term premium and expectation errors about the short-term rate, and the extent of interest rates and bond returns predictability in a model of drifting yields, all warrant further research.

This article shows that reconstructing the term structure starting from a simple monetary policy rule with an equilibrium rate driven by productivity and demographics trends, together with long-term inflation expectations, goes a long way in capturing the stochastic trend in yields. Our framework establishes a set of novel facts about Treasury bonds, while offering the possibility to revisit classic questions related to bond predictability and monetary inertia.

First, our monetary policy rule (with a target rate modeled by fluctuations in potential output, demographics, and long-term inflation expectations) tracks well the evolution of the short-term rate both in- and out-of-sample. Importantly, by being explicit about the nonstationary drivers of rates, our model is purposely transparent and simple (i.e., not involving any filtering). We find that policy inertia can be overestimated if the drivers of the drifting equilibrium policy rates are not included in the monetary reaction function, contributing to the debate on monetary policy inertia as a result of omitted factors in the Fed's reaction function or interest rate smoothing (see, e.g., Rudebusch (2002), (2006), Coibion and Gorodnichenko (2012)).

Second, we derive the implications of our monetary policy rule specification for the entire term structure of Treasury bond yields. Our approach decomposes bond yields at any maturity into a drifting component (the average expected sequence of monetary policy rates over the life of the bond) and a residual cyclical component (the deviation of yields from their drift). We show that our framework with drifting bond prices implies a battery of misspecification tests such as parametric restrictions on yields and their drift that are analogous to the restriction between prices and dividends in the Campbell and Shiller (1988) present-value

[2]The relevance of investigating the drift in the term structure of yields is not restricted to Treasury bonds. For example, Farhi and Gourio (2018) propose a macro-finance neoclassical growth model to account for drifting real rates and stable return to private capital. van Binsbergen (2020) finds that accounting for secular trends in interest rates is fundamental for assessing long-duration dividend risk. Campbell and Sigalov (2020) derive a model of reaching for yield and show that agents take more risk when the real interest rate declines while the risk premium remains constant. Also, see Campbell (2019) for a discussion (available at https://scholar.harvard.edu/files/campbell/files/nber_ltamkeynoteslides.pdf) on the importance of drifting prices for long-term investing.

model. Specifically, when the (nonstationary) drivers of the monetary policy rates have been correctly specified, deviations of bond prices from their estimated drift should be stationary with a cointegrating vector of $(1, -1)$, generating the cyclical components of yields. Our empirical analysis confirms these predictions.

Having analyzed the statistical properties of our model, and confirmed it is well-behaved, we investigate bond risk premia predictability within our framework with drifting bond prices. We formally show that stationary deviations of bond prices from their drift should predict excess bond returns. Empirically, our model generates large R^2 of about 30% (10%) when it is used to predict the 1-year (1-quarter) ahead excess returns on bond with maturities ranging from 2 to 10 years. We also construct a single yield-based cycle factor and find that our return-forecasting factor subsumes common bond risk premia predictors, such as the Cochrane and Piazzesi (2005) and Cieslak and Povala (2015) factors. Importantly, our results survive out-of-sample and hold internationally.

In the last part of the article, we address the, admittedly challenging, question of the economic interpretation of the cyclical components. Within our stylized framework, we supply an upper bound to the role of deviations from rational expectations (RE) (in the form of diagnostic expectations (DE)) for the fluctuations in the cyclical component of yields. In particular, when we test for the role of DE (overreaction of agents to deviations of the monetary policy rate from its trend), we find that up to 40% of the fluctuations in yield cycles can indeed be attributed to this mechanism for bonds with a 2-year maturity. However, the explanatory power of DE for bond cycles declines with the maturity of the bond, leaving an important role for term premia.

Related Literature

Our evidence that bond prices are drifting is in line with several papers documenting a slow-moving component common to the entire term structure (see, e.g., Balduzzi, Das, and Foresi (1998), Fama (2006)). Although a factor model for bond yields can admit a unit root in the feedback autoregressive matrix, OLS estimates of near-unit roots are notoriously biased downward, thus overestimating the amount of mean reversion in yields.[3] To address this issue, an important and growing literature has modeled Treasury yields using shifting endpoints (Kozicki and Tinsley (2001)), near-cointegration (Jardet, Monfort, and Pegoraro (2013)) or long memory (Golinski and Zaffaroni (2016)), vector autoregressive models (VAR) with common trends (Negro, Giannone, Giannoni, and Tambalotti (2017)), slow-moving averages of inflation (Cieslak and Povala (2015)) and consumption (Jørgensen (2018)), or an (unobserved) stochastic trend common across Treasury yields (Bauer and Rudebusch (2020)). We contribute to this important literature by proposing a cohesive (cointegrated) framework with observable economic trends to explore the implications of drifting equilibrium rates for monetary policy, Treasury yields, and bond returns predictability.[4]

[3]Piazzesi and Cochrane (2008) propose to fit the cross section via an affine model to reduce some of the statistical uncertainty surrounding level stationarity versus unit roots specifications.

[4]Also, in our framework stationarity of bond returns naturally coexists with nonstationary bond prices. Bond returns are predicted by the stationary deviations of bond prices from their drift.

Jordà and Taylor (2019) and Feunou and Fontaine (2023) couple simple restrictions from New Keynesian IS and Phillips curves with a cyclical and secular decomposition of short rates and output to study monetary policy divergence and the impact of nominal shocks, respectively. We complement this line of research by documenting the importance of a transitory-permanent decomposition of bond yields for interest rates and bond return predictions. Furthermore, we provide evidence that demographics together with the growth rate of potential output are important drivers of the secular trend in bond yields.

Finally, our article fits into the literature that studies the role played by (shifts in) the monetary conduct in determining the dynamics of bond yields.[5] Berardi, Markovich, Plazzi, and Tamoni (2020) show that the stance of monetary policy (as proxied by the difference between the natural rate of interest and the current level of short-term rate) contains valuable information for bond predictability. Ang, Boivin, Dong, and Loo-Kung (2011) show that the evolution of the Fed's response to inflation affects long-term yields. Similarly to Ang et al. (2011), we propose to model monetary policy and the term structure of interest rates jointly. However, our modeling of the policy rule with a drifting equilibrium rate is different from their model with time-varying policy coefficients. In turn, our approach has implications for interest rates comovement and bond returns predictability induced by deviations of bond prices from their drift. These testable implications are unique to our framework and not shared by Ang et al. (2011).

II. Modeling Monetary Policy

Monetary policy rules specify the dynamics of the short-term rate, $y_t^{(1)}$. The following specification is general and encompasses most of the rules that have been proposed in the literature:

$$y_t^{(1)} = y_t^* + \beta' X_t + u_t^{(1)}, \tag{1}$$

where y_t^* is the equilibrium monetary policy rate,[6] X_t is a vector of stationary monetary policy factors, and $u_t^{(1)}$ is a monetary policy residual following an AR(1) process with persistence ρ (e.g., Rudebusch (2006) and Pasten, Schoenle, and

Interestingly Bauer and Rudebusch (2020) note that, even when no arbitrage is imposed, the loading of returns on the unobserved common stochastic trend is an order of magnitude smaller than the loading of prices. They also report that predictive regressions of yields on detrended yields and trend proxies lead to coefficients on the trend that are not significantly different from 0.

[5]An important literature (see, e.g., Bernanke and Kuttner (2005), Ozdagli (2018), and Chava and Hsu (2020)) investigate the impact of monetary policy shocks on equity prices and the cross section of stock returns. Koijen, Lustig, and Van Nieuwerburgh (2017) propose a 23-factor model for stocks and bond returns. The investigation of a factor model with drifting bond *and equity prices is an interesting avenue for future research.*

[6]The "natural" level of real interest rates is often referred to as the "natural," "equilibrium," or "neutral" real rate of interest. Interestingly, the possibility of a nonstationary equilibrium rate is rarely entertained in the traditional literature. A notable exception is Woodford (2001) who shows that the optimal policy response to real disturbances requires including a time-varying real rate in monetary policy rules. See Giammarioli and Valla (2004) and Kiley (2015) for a review of the various concepts and estimation methods adopted in the literature.

Weber (2020)). Arguably, the most famous special case of this specification is the Taylor (1993) rule. In this case, the vector X_t is composed of the output gap and the percentage deviation of inflation from its target. Furthermore, the Taylor (1993) rule assumes a constant equilibrium policy rate (i.e., $y_t^* = y^*$) and provides a natural benchmark for our analysis.

With a constant equilibrium rate, the estimate of the AR(1) persistence parameter is often close to one. This is to be expected since, if monetary policy rates are drifting, any attempt to model them only by means of stationary factors such as the output and inflation gaps naturally leads to a highly persistent process for $u_t^{(1)}$. This fact has spurred an important literature debating the sources and the implications of monetary policy inertia.[7] One common narrative is that monetary policy inertia is fictitious and stems from omitted variables in the Fed's reaction function (e.g., Rudebusch (2002), (2006)). Others have argued that the central bank conducts sluggish partial adjustment of short-term policy interest rates, modeled through interest smoothing in the policy rule (e.g., Coibion and Gorodnichenko (2012)).

We contribute to this debate showing that monetary policy inertia is overestimated when the time-varying drivers of the drifting equilibrium rate are not included in the monetary policy rule.

Interest rates are sometimes modeled in first-difference which removes the stochastic trend in policy rate at the cost of leaving the equilibrium level of the policy rate undetermined (e.g., Orphanides (2003)). The model in first-difference is a special case of our general specification when $\rho = 1$. Specifying the monetary policy rule in first-difference comes with benefits and costs.[8] The benefit of making the rule independent from the challenging estimation of the level of the equilibrium rate has to be traded-off against the cost of accepting that any monetary policy shock (i.e., any deviation from the rule) has a permanent effect on policy rates. Indeterminacy is a major concern for long-term forecasting, because as the unconditional distribution of policy rates is not defined, the long-run policy rate is also left undetermined.

We propose a "cointegrating" approach to drifting policy rates, where the stationarity of residuals of the monetary policy reaction function is taken as an indication of a valid specification for y_t^*. Equivalently, a valid specification for the equilibrium rate requires that y_t^* is the stochastic trend that drives drifting policy rates.[9]

Specifically, we propose to model drifting policy rates as follows:

(2)
$$\begin{aligned} y_t^{(1)} &= y_t^* + \beta_1 E_t\left(\pi_{t+1} - \pi_{t+1}^*\right) + \beta_2 E_t(x_{t+1}) + u_t^{(1)} \\ y_t^* &= \gamma_1 \mathrm{MY}_t + \gamma_2 \Delta x_t^{\mathrm{pot}} + \gamma_3 \pi_t^* \\ u_t^{(1)} &= \rho u_{t-1}^{(1)} + \varepsilon_t^{(1)}, \end{aligned}$$

[7]For a detailed discussion on optimal monetary policy inertia see Woodford ((2001), (2003)).

[8]Cochrane (2007) provides a thorough discussion on the effects of specifying a model in level versus first-difference to compute long-term yield-curve decomposition.

[9]We assume the component y_t^* of the short-rate to be integrated of order one (i.e., $I(1)$) and model it using demographics and productivity trends, plus long-term inflation expectations. This is an over-identifying restriction in the sense that one could make other assumptions (e.g., $y_t^* \sim I(d)$). We thank the referee for this observation.

where $y_t^{(1)}$ is the one-period (23-month) yield, y_t^* is the equilibrium nominal rate, π_t is the percentage annual log change in personal consumption expenditures (PCE), π_t^* is the Fed perceived target rate (PTR), and x_t is the output gap (log percentage difference between real GDP and potential GDP). The monetary policy residual with a drifting equilibrium rate $u_t^{(1)}$ is stationary under cointegration (i.e., $|\rho| < 1$) and $\varepsilon_t^{(1)}$ is an IID innovation. The drivers of the equilibrium real rate are the age structure of population (MY) and potential output growth (Δx_t^{pot}).[10] We obtain the nominal equilibrium rate by adding the central bank inflation target π_t^*. The Appendix provides details on the data source.

Our choice for potential output growth and demographics as drivers of the equilibrium real rate borrows from Laubach and Williams (2003) and Jordà and Taylor (2019). In particular, Jordà and Taylor (2019) show that the decline of the neutral rate cannot be explained just by the decline in the growth rate of potential. We link this unobserved component to demographic. Specifically, as demographic variable we use the ratio of middle-aged (40–49) to young (20–29) population in the United States (labeled as MY). The use of this variable is motivated by the overlapping generation model of Geanakoplos, Magill, and Quinzii (2002) which predicts a negative relation between Treasury yields and MY. Potential output growth is the percentage annual log change in potential output. MY_t, Δx_t^{pot}, and π_t^* are nonstationary (i.e., they are integrated of order 1) and they represent the drivers of the drifting equilibrium rate in our cointegrated specification.[11,12]

Finally, in our tests, we always compare the results from our baseline (drifting) model to the results of a restricted model that, inspired by the large body of literature on the classical Taylor (1993) rule, does not model the drift in monetary policy[13]:

---

[10]Including only inflation as driver of nonstationary policy rates is equivalent to assume a counterfactual stationary equilibrium real rate (see, e.g., Lunsford and West (2019)). Empirically, although inflation is the most important driver of the policy rate, using only inflation leaves a persistent component in yields unexplained (i.e., the AR(1) persistence parameter for $u_t^{(1)}$ in equation (2) is 0.79); this evidence is in line with Bauer and Rudebusch (2020).

[11]Our specification is compatible with yields being nonstationary or yields appearing nonstationary from the perspective of a model that does not include regime shifts. What matters for the validity of our specification is that the deviations of actual rates from equilibrium rates are stationary.

[12]The p-values from the Phillips and Perron (1988) unit root test for MY_t, Δx_t^{pot}, and π_t^* are respectively 0.95, 0.13, and 0.69; thus, we cannot reject the null of each series being integrated of order 1. Furthermore, since the seminal work of Laubach and Williams (2003), it is standard to take that the growth of potential output as integrated of order 1.

[13]We deviate in two respects from a standard empirical Taylor rule. First, the model in (3) is a forward-looking version of the policy rule. This is consistent with the perspective that monetary policy changes take time to affect the economy (see, e.g., Clarida, Gali, and Gertler (2000), Coibion and Gorodnichenko (2011)). Second, we specify the inflation gap as deviations of inflation from a time-varying inflation target (π_t^*) rather than from a constant inflation target (e.g., 2%). This is in line with the idea that the inflation gap should measure the difference between actual inflation and the central bank's long-run target (e.g., Cogley, Primiceri, and Sargent (2010)). A standard empirical Taylor rule with constant inflation target would lead to a better in-sample fit ($R^2 = 66\%$). The in-sample success of the standard Taylor rule is explained by the fact that detrending inflation with a constant target over the full sample results in a highly persistent inflation gap, which mechanically explains better nonstationary yields. Importantly, a standard Taylor rule is inferior to our rule with drifting rates in terms of long-term forecasts, modeling the term structure, and bond risk premia predictability.

(3) $$y_t^{(1)} = y^* + \beta_1 E_t\left(\pi_{t+1} - \pi_{t+1}^*\right) + \beta_2 E_t(x_{t+1}) + u_t^{(1)}.$$
$$u_t^{(1)} = \rho u_{t-1}^{(1)} + \varepsilon_t^{(1)}.$$

Empirical Results

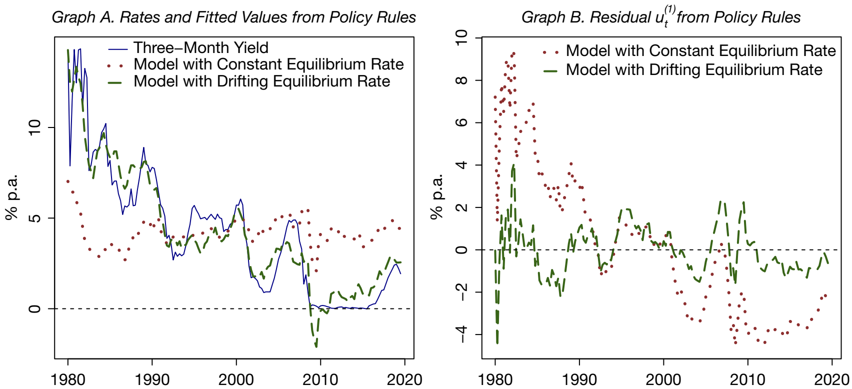
Graph A of Figure 1 displays the realized nominal short-term rate, the fitted rates from our cointegrated monetary rule (cf. equation (2)), and the fitted monetary policy rates from a version of our model which restricts the equilibrium rate to be constant (cf. equation (3)). Graph B plots the monetary policy residuals implied by our proposed monetary policy rule and its restricted version. Table 1 reports the estimation results for these two rules.[14]

Graph A of Figure 1 shows that our monetary rule with a drifting equilibrium rate tracks well the short-term rate movements throughout the sample. Indeed, the R^2 for the cointegrated specification is about 95% whereas that of a model with constant equilibrium rate is just about 4% (cf. Table 1).[15,16] Graph B of Figure 1

FIGURE 1

Actual Versus Fitted Short-Term Rate

Graph A of Figure 1 shows the actual 3-month yield and fitted values for our (cointegrated) model with drifting equilibrium rates (cf equation (2); see green dashed line) as well as for a model that restricts the equilibrium rate to be constant (cf equation (3); see brown dotted line). Graph B shows the differences between actual 3-months yield and the fitted values. Quarterly observations. The sample period is 1980:Q1 to 2019:Q4.



[14]Our estimate of the loading on π_t^* is in line with parameter values reported in Bauer and Rudebusch ((2020), Table 1) despite the difference in the maturity of the bond analyzed (their Table 1 analyzes the 10-year bond, whereas we focus on the 3-month Treasury bill).

[15]Furthermore, a regression of the 23-month yield on the fitted values implied by the two monetary rules (dotted and dashed lines in Graph A of Figure 1) delivers an estimate of 0 on the rule with constant equilibrium rates (3), and a statistically significant estimate not different from one on the drifting rule (2).

[16]Positing the following cointegration framework where the equilibrium real rate r_t^* is estimated first, that is,

$$y_t^{(1)} = \alpha_1 r_t^* + \alpha_2 \pi_t^* + \beta_1 E_t\left(\pi_{t+1} - \pi_{t+1}^*\right) + \beta_2 E_t(x_{t+1}) + u_t^{(1)}$$
$$r_t^* = \gamma_1 \text{MY}_t + \gamma_2 \Delta x_t^{\text{pot}},$$

leaves our conclusions unaltered. See Figure A.1 in the Supplementary Material.

TABLE 1

Short-Term Rate Models With and Without Drifting Equilibrium Rate

Table 1 reports the estimates for our (cointegrated) model with drifting equilibrium rates (cf equation (2); see column 2) as well as estimates for a model that restricts the equilibrium rate to be constant (cf equation (3); see column 1). We estimate the two rules by instrumental variables, where the instruments are lags of inflation gap and output gap. The last row reports OLS estimates for the monetary policy residuals' persistence. Values in parentheses are GMM standard errors that correct for autocorrelation in the residuals. Quarterly observations. The sample period is 1980:Q1 to 2019:Q4. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

	Three-Month Yield	
	1	2
MY		−2.652***
		(0.726)
Δx_t^{pot}		0.932***
		(0.317)
π_t^*		1.656***
		(0.177)
$E_t(\pi_{t+1} - \pi_{t+1}^*)$	0.721	0.709***
	(0.519)	(0.244)
$E_t(x_{t+1})$	0.086	0.389***
	(0.481)	(0.137)
Constant	4.656***	
	(1.036)	
No. of obs.	160	160
Adj. R^2	0.036	0.950
ρ	0.949***	0.673***
	(0.022)	(0.110)

shows that the residuals implied by our drifting monetary policy rule are mean reverting. On the other hand, the residuals from a rule with constant equilibrium rates display a close-to-unit root behavior. This is confirmed in Table 1: The residuals from the rule with drifting (constant) equilibrium rates have an autoregressive coefficient equal to 0.67 (0.95).

Figure 2 displays the forecasts implied by the two monetary policy rules. Although the in-sample performance from the two models is similar, the long-term out-of-sample forecasts are different. Indeed, the policy rule with constant equilibrium rates generates forecasts that converge fast to the unconditional mean.[17] On the other hand, the drifting monetary policy rule tracks well the future evolution of the short rate for each of the three out-of-sample periods considered in the figure. Figures A.2 and A.3 of the Supplementary Material confirm that allowing for interest rate smoothing in the rule with constant equilibrium rate does not alter our conclusion: The long-term forecasts converge fast to the unconditional mean, and underperform relative to the forecasts from a model with drifting equilibrium rate. This conclusion holds independently from whether interest smoothing is characterized as a first- or a second-order autoregressive process (Coibion and Gorodnichenko (2011)).

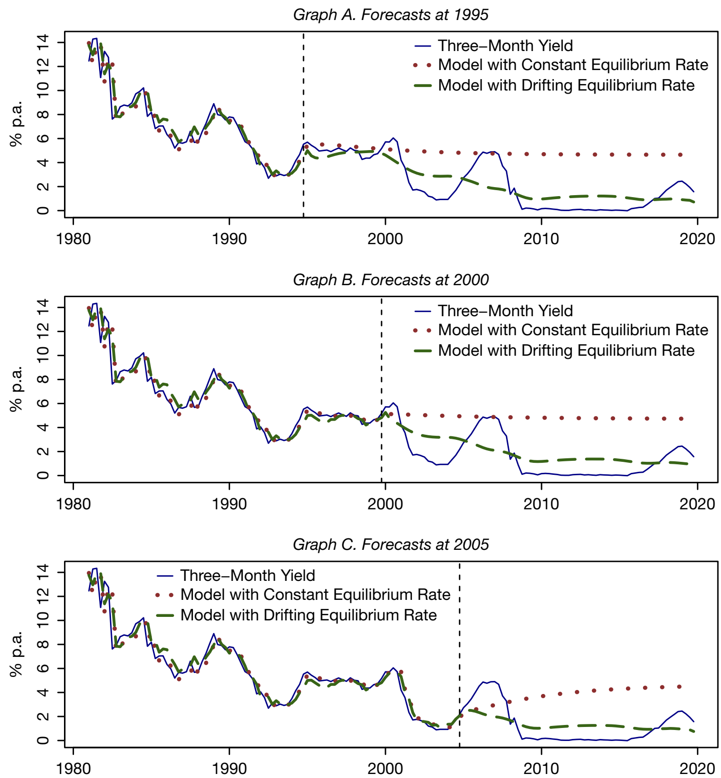
Finally, note that the fitted short rate in Figure 2 falls below 0 only for a very short period of time, and the forecasts in Figure 2 never hit the bound. As the economy is affected by the entire path of expected future short-term rates (e.g., Swanson and Williams (2014)), our results suggest that an accurate modeling of the

[17]Rudebusch (2002) highlights the tension between the apparent high persistence and low predictability of policy rates.

FIGURE 2

Long-Term Forecasts of Short-Term Rate

Figure 2 shows the actual 3-month yield and predicted rates implied by our (cointegrated) model with drifting equilibrium rates (cf equation (2); green dashed line) and by a model that restricts the equilibrium rate to be constant (cf equation (3); brown dotted line). The forecast of the drifting rule exploits the exogeneity of the demographic variable (MY) and of potential output (Δx^{pot}). In particular, the rule is estimated until 1995, 2000, and 2005 in Graphs A, B, and C, respectively. We then use the coefficients estimates, the projections of MY and Δx^{pot} (see also the Appendix), and the forecast of inflation and output gap from a VAR(1) as in equations (7) and (8). π^* is modeled as a random walk. Monetary policy residuals persistence ρ is 0.673 and 0.949 for the drifting and the constant equilibrium rate models respectively. Dotted vertical lines represent the end of in-sample estimation period. Quarterly observations. The sample period is 1980:Q1 to 2019:Q4.



trend alleviates concerns related to the effectiveness of monetary policy at the zero lower bound (see Aruoba and Schorfheide (2016) for a discussion).

In all, our evidence points to the importance of modeling the economic determinants of the time-varying equilibrium rates. This is consistent with Jordà and Taylor (2019) who document that a policy rule with a time-invariant intercept (as in equation (3)) is rejected in international data.

III. Modeling a Drifting Term Structure

The entire term structure is drifting. Models that parsimoniously describe the term structure by projecting rates on a set of factors and by modeling the dynamics of the factors with a VAR will be inevitably confronted with the problem generated by the presence of unit roots in the VAR. Highly persistent VAR generate imprecise forecasts at long horizons (e.g., Giannone, Lenza, and Primiceri (2019)). This

https://doi.org/10.1017/S0022109022001557 Published online by Cambridge University Press

feature can explain mixed results from the forecasting performance of affine term structure models (see, e.g., Duffee (2002), Sarno, Schneider, and Wagner (2016)). We propose to use the drift in monetary policy rates to model the drift in the entire term structure:

$$
\begin{aligned}
y_t^{(n)} &= y_t^{(n),*} + \delta_0^{(n)} + u_t^{(n)}, \\
y_t^{(n),*} &= \left(\frac{1}{n}\right) \sum_{i=0}^{n-1} E_t\left[y_{t+i}^{(1)}\right], \\
u_t^{(n)} &= \rho^{(n)} u_{t-1}^{(n)} + \varepsilon_t^{(n)}.
\end{aligned} \tag{4}
$$

Yields at all maturities are decomposed into a trend, $y_t^{(n),*}$, and a cyclical component, $\delta_0^{(n)} + u_t^{(n)}$. The trend is the average of expected monetary policy rates over the duration of the bond, while the cyclical component is the stationary residuals from the $(1,-1)$ cointegrating relationship between yields and their drift. We consider as valid any model of the term structure that delivers cointegration between $y_t^{(n)}$ and $y_t^{(n),*}$ with a $(1,-1)$ cointegrating vector and, therefore, a stationary $u_t^{(n)}$.

Our full-term structure model is specified as follows:

$$
\begin{aligned}
y_t^{(1)} &= y_t^* + \beta_1 E_t\left(\pi_{t+1} - \pi_{t+1}^*\right) + \beta_2 E_t(x_{t+1}) + u_t^{(1)}, \\
y_t^* &= \gamma_1 \mathrm{MY}_t + \gamma_2 \Delta x_t^{\mathrm{pot}} + \gamma_3 \pi_t^*, \\
u_t^{(1)} &= \rho u_{t-1}^{(1)} + \varepsilon_t^{(1)},
\end{aligned} \tag{5}
$$

$$
\begin{aligned}
y_t^{(n)} &= y_t^{(n),*} + \delta_0^{(n)} + u_t^{(n)}, \\
u_t^{(n)} &= \rho^{(n)} u_{t-1}^{(n)} + \varepsilon_t^{(n)}, \\
y_t^{(n),*} &= \left(\frac{1}{n}\right) \sum_{i=0}^{n-1} E_t\left[y_{t+i}^{(1)}\right],
\end{aligned} \tag{6}
$$

$$
\left(\pi_t - \pi_t^*\right) = \theta_{1,1}\left(\pi_{t-1} - \pi_{t-1}^*\right) + \theta_{1,2} x_{t-1} + \theta_{1,3}\left(y_{t-1}^{(1)} - y_{t-1}^*\right) + v_{1,t}, \tag{7}
$$

$$
x_t = \theta_{2,1}\left(\pi_{t-1} - \pi_{t-1}^*\right) + \theta_{2,2} x_{t-1} + \theta_{2,3}\left(y_{t-1}^{(1)} - y_{t-1}^*\right) + v_{2,t}, \tag{8}
$$

where we assume $\mathrm{cov}\left(v_{1,t}, u_t^{(1)}\right) = \mathrm{cov}\left(v_{2,t}, u_t^{(1)}\right) = 0$.

Projections of the equilibrium policy rates depend on productivity and demographics, which we take as exogenous; thus, we do not specify the law of motion for MY_t and $\Delta x_t^{\mathrm{pot}}$). The U.S. Census Bureau and the U.S. Congressional Budget Office provide ready-to-use projections respectively for MY and potential output. Equations (7) and (8) are used to compute the projections of inflation and output gaps. The dynamics of these two stationary variables depend on their own lags and on a third stationary variable: the deviation of the short-term rate from its trend. This cycle in monetary policy enters the dynamics of output and inflation gaps with a 1-quarter lag; this is consistent with the delay with which monetary policy affects these variables in our specification of the forward-looking policy rule (5). Finally, note that we do not impose no-arbitrage (NA) restrictions when estimating our model. Thus, our estimation strategy runs the cost of losing efficiency if NA holds to gain consistency in the case NA is violated.

A. Empirical Results

1. Misspecification Test for Term Structure Models

The validity of a model with drifting monetary policy rates and bond prices can be assessed by checking the existence of cointegrating relationships with parameters $(1,-1)$ between $y_t^{(n)}$ and $y_t^{(n),*}$ (see equation (6)). Thus, in this section, we investigate the strength of the cointegrating relationship, the $(1,-1)$ parametric restriction, and the behavior of the residuals for our baseline model (see equations (5)–(8)) as well as for its restricted version where the drift in monetary policy is assumed away (i.e., $y_t^* = y^*$).

Figure 3 reports the results for the cointegration relationship for five maturities ranging from 2 ($n = 8$ quarters) to 10 years ($n = 40$ quarters). Graph A is for the restricted model, whereas Graph B is for our model with drifting equilibrium policy rates.

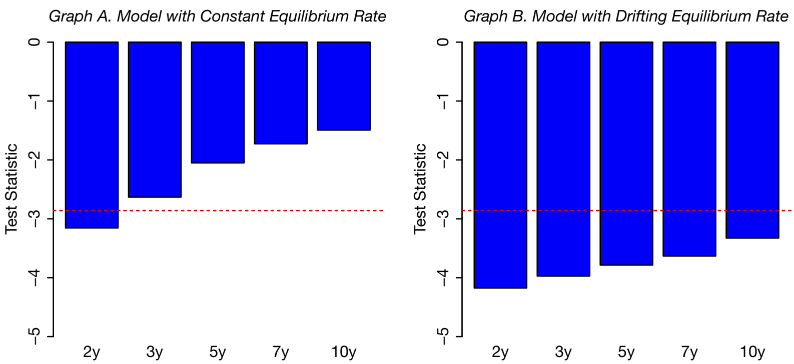
Our model with drifting equilibrium monetary policy rate provides overwhelming evidence to reject the null hypothesis of absence of cointegrating relation between $y_t^{(n)}$ and $y_t^{(n),*}$ for all the considered maturities. From an economic perspective, this implies that fluctuations in productivity, demographics, and long-term inflation expectations are successful in modeling not only the drift in monetary policy rates but also the drift in the entire term structure.

Furthermore, Table A.1 of the Supplementary Material confirms that, within our framework with drifting policy rates, the parametric restriction $(1,-1)$ on the cointegrating relationship between yields and their drift is supported in the data for every maturities ranging from 2 to 10 years.

FIGURE 3

Engle and Granger (1987) Cointegration Test

Figure 3 shows results for the Engle and Granger (1987) cointegration test for $u_t^{(n)}$ defined in equation (4) for different maturities. Graph A reports test statistics for a model that restricts the equilibrium policy rate to be constant, that is, $y_t^* = y^*$ (cf equation (3)). Graph B reports test statistics for our (cointegrated) model with drifting equilibrium policy rates (cf equations (5)–(8)). The null hypothesis is absence of cointegration. The dashed red line is the critical value at 5% level of significance as suggested by MacKinnon (2010). Quarterly observations. The sample period is 1980:Q1 to 2019:Q4.



In all, our choice of the drivers for the equilibrium rate y_t^* provides also an accurate description of the stochastic trend underlying interest rates.

B. Predicting Holding Period Excess Returns

Predictability of interest rates on the basis of a model with a common stochastic trend in yields also implies predictability of holding period excess returns on the basis of the stationary deviations of bond yields from their drift.

To see this, write the expected excess return obtained by holding the n-period bond for one period as:

$$\begin{aligned}(9)\quad E_t\left(rx_{t+1}^{(n)}\right) &= y_t^{(n)}n-(n-1)E_t\left(y_{t+1}^{(n-1)}\right)-y_t^{(1)}\\ &= y_t^{(n)}-(n-1)\left(E_t\left(y_{t+1}^{(n-1)}\right)-y_t^{(n)}\right)-y_t^{(1)}\\ &= y_t^{(n)}-y_t^{(1)}-(n-1)\left(E_t\left(y_{t+1}^{(n-1)}\right)-y_t^{(n-1)}\right)-(n-1)\left(y_t^{(n-1)}-y_t^{(n)}\right),\end{aligned}$$

where $y_t^{(n)}-y_t^{(1)}$ is the slope of the term structure, $\left(y_t^{(n-1)}-y_t^{(n)}\right)$ is known as the roll-down, and $\left(E_t\left(y_{t+1}^{(n-1)}\right)-y_t^{(n-1)}\right)$ is the expected change in prices of the $(n-1)$-maturity bond. Since the seminal contributions by Fama and Bliss (1987) and Campbell and Shiller (1991), the slope of the term structure has played a central role in forecasting bond returns. Indeed, it is common to assume away any predictability arising from $\left(E_t\left(y_{t+1}^{(n-1)}\right)-y_t^{(n-1)}\right)$, since the level of the term structure is deemed to be close to unforecastable (see, e.g., Duffee (2013)).

Our "cointegrated" specification of the monetary policy rule and the term structure suggests otherwise. Using equation (6) one can express the expected price changes as

$$(10)\quad E_t\left(y_{t+1}^{(n-1)}\right)-y_t^{(n-1)}=E_t\left(y_{t+1}^{(n-1),*}-y_t^{(n-1),*}\right)+\left(\rho_{(n-1)}-1\right)\underbrace{\left(y_t^{(n-1)}-y_t^{(n-1),*}-\delta_0\right)}_{u_t^{(n-1)}}.$$

Therefore, in our model, persistent but stationary deviations of bond prices from their drift, $u_t^{(n-1)}$, show up as a natural predictor of excess bond returns.[18] This term has gone unrecognized since standard models start off with stationary factor (within our framework, this is equivalent to assume a constant equilibrium rate). In turn, this leads to a close-to-unit-root residual (cf. Graph A of Figure 4), or $\rho_{(n-1)}-1\approx 0$ $\left(\text{and the level being a random walk } E_t\left(y_{t+1}^{(n-1)}\right)=y_t^{(n-1)}\right)$.[19]

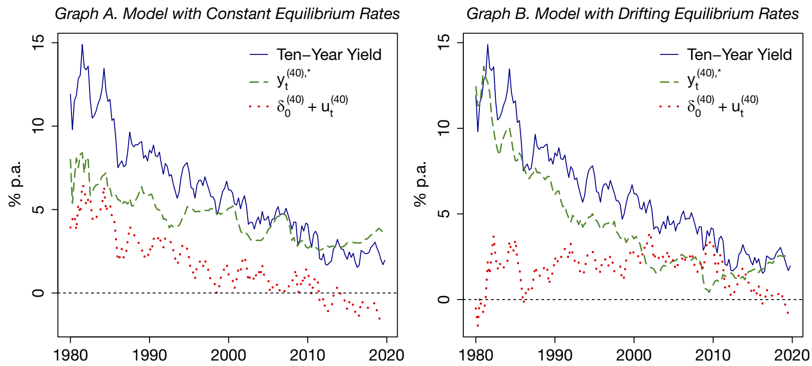
[18]More precisely, $u_t^{(n-1)}$ should forecast the price change component in bond returns. However, empirically the correlation between $rx_{t+4}^{(n)}$ and the price change term, $-\left(y_{t+4}^{(n-4)}-y_t^{(n-4)}\right)$, is high at 93%, 95%, 97%, 98%, and 99% for $n=8,12,20,28,40$ quarters, respectively.

[19]Cieslak and Povala (2015) and Jørgensen (2018) predict bond returns using a detrended (term structure) level factor. Using their proposed persistence-based Wold decomposition, Ortu, Severino, Tamoni, and Tebaldi (2020) extract a cyclical component from the level of the yield curve and show that it contains information about future excess bond returns. To our knowledge, we are the first to show that a cyclical component of the level of the term structure emerges as a natural predictor within a cointegrated framework of bond prices.

FIGURE 4

Decomposing Long-Term Rates

Graph A of Figure 4 shows the decomposition of the 10-year yield implied by a model which assumes away drifting monetary policy rates (i.e., $y_t^* = y^*$). Graph B shows the decomposition of the 10-year yield implied by our model with drifting equilibrium rates (see equations (5)–(8)). Quarterly observations. The sample period is 1980:Q1 to 2019:Q4.



We start the evaluation of the predictive performance of our model with a drifting equilibrium rate by running the following regression:

$$rx_{t+4}^{(n)} = \alpha + \beta E_t\left(rx_{t+4}^{(n)}\right) + \varepsilon_t, \tag{11}$$

where $rx_{t+4}^{(n)}$ is the realized 1-year holding period excess return of a bond with maturity n-quarters. We denote with $E_t\left(rx_{t+4}^{(n)}\right)$ the expected excess return implied by our specification that allows for stationary deviations of bond prices from their drifts.[20] We compare our specification to the classical model with a constant equilibrium rate.[21] Table 2 displays the results for the model with a constant equilibrium rate in Panel A, and the results for our model with drifting bond prices in Panel B. We consider maturities ranging from 2 ($n = 8$ quarters) to 10 years ($n = 40$ quarters). The regression of realized excess returns on the expected returns implied by our (cointegrated) model with drifting equilibrium rates delivers statistically significant estimates and coefficients of determination that are greater than 30% at all maturities.[22] On the other hand, a classical model with constant equilibrium rates leads to a coefficient not significantly different from 0 and to small explanatory power.

We also highlight that the model with constant equilibrium rates performs worse than a (reduced-form) model based just on the slope. This is easily explained.

[20]We exploit equations (5)–(8) together with the exogeneity of demographics and potential output to construct the expected change in constant-maturity yield $\left(E_t\left(y_{t+1}^{(n-1)}\right) - y_t^{(n-1)}\right)$ in equation (9).

[21]To make our results comparable to a large literature (e.g., Cochrane and Piazzesi (2005), Cieslak and Povala (2015)) we focus on 1-year excess returns. However, our conclusions are identical when we use 1-quarter holding period returns.

[22]The constant is not statistically significant for bond with maturities $n = 8,12,20$ quarters. Panel A of Figure A.4 in the Supplementary Material displays the realized and fitted values for bonds with 2- and 10-years maturity.

TABLE 2

Predictive Regressions Across Different Maturities

Table 2 reports OLS estimates for the regression $rx_{t+4}^{(n)} = \alpha + \beta E_t\left(rx_{t+4}^{(n)}\right) + \varepsilon_t$, where $rx_{t+4}^{(n)}$ is the realized 1-year holding period excess return of a bond with maturity n-period and $E_t\left(rx_{t+4}^{(n)}\right)$ is the expected excess return implied by our specifications. Panel A reports results for the classical model with a constant equilibrium rate. Panel B reports results for our model with drifting equilibrium rates. Values in parentheses are conservative standard errors from reverse regressions computed as in Hodrick (1992). Constant estimates are not tabulated. Quarterly observations. The sample period is 1980:Q1 to 2019:Q4. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

	$rx_{t+4}^{(8)}$	$rx_{t+4}^{(12)}$	$rx_{t+4}^{(20)}$	$rx_{t+4}^{(28)}$	$rx_{t+4}^{(40)}$
	1	2	3	4	5
Panel A. Model with Constant Equilibrium Rate					
$E_t\left(rx_{t+4}^{(8)}\right)$	0.427**				
	(0.201)				
$E_t\left(rx_{t+4}^{(12)}\right)$		0.308			
		(0.188)			
$E_t\left(rx_{t+4}^{(20)}\right)$			0.212		
			(0.167)		
$E_t\left(rx_{t+4}^{(28)}\right)$				0.171	
				(0.151)	
$E_t\left(rx_{t+4}^{(40)}\right)$					0.139
					(0.131)
No. of obs.	156	156	156	156	156
R^2	0.128	0.092	0.069	0.060	0.054
Panel B. Model with Drifting Equilibrium Rate					
$E_t\left(rx_{t+4}^{(8)}\right)$	0.868***				
	(0.187)				
$E_t\left(rx_{t+4}^{(12)}\right)$		0.822***			
		(0.159)			
$E_t\left(rx_{t+4}^{(20)}\right)$			0.693***		
			(0.151)		
$E_t\left(rx_{t+4}^{(28)}\right)$				0.707***	
				(0.120)	
$E_t\left(rx_{t+4}^{(40)}\right)$					0.619***
					(0.120)
No. of obs.	156	156	156	156	156
Adj. R^2	0.366	0.366	0.331	0.361	0.320

The realized returns $rx_{t+4}^{(n)}$ on the left-hand side of (11) are stationary, whereas the expected returns $E_t\left(rx_{t+4}^{(n)}\right)$ from the model with constant equilibrium rate is nonstationary since it inherits the drift from the residual component $u_t^{(n)}$ (cf. Figure 4).

1. Dissecting Predictive Regressions

To further dissect the unique contribution coming from our cointegrated approach, Table 3 shows that the expected change in the $(n-1)$-maturity bond prices drives away the predictability of the slope (column 1), and that deviations of bond prices from their drift, $u_t^{(n-1)}$, are the most important driver of such predictability (cf. columns 3 and 4). Also, the loading on the cyclical component $u_t^{(n-1)}$ is negative as predicted by our framework: If $0 < \rho_{(n-1)} < 1$, then next period returns are negative in times when bond prices are higher than those implied by their drift.

In the Supplementary Material, we show that the relevance of such cyclical component for forecasting excess returns is not restricted to any specific maturity or holding period. Table A.2 of the Supplementary Material reports results for the

TABLE 3

Dissecting Predictive Regressions

Table 3 reports OLS estimates for the regression $rx_{t+4}^{(40)} = \alpha + \beta' X_t + \varepsilon_t$, where $rx_{t+4}^{(40)}$ is the realized 1-year holding period excess return of a bond with maturity 10-year and X_t contains different return predictors. Column 1 exploits equation (9) reported here for the reader's convenience:

$$E_t\left(rx_{t+4}^{(40)}\right) = y_t^{(40)} - y_t^{(4)} - (40-4)\left(E_t\left(y_{t+4}^{(40-4)}\right) - y_t^{(40-4)}\right) - (40-4)\left(y_t^{(40-4)} - y_t^{(40)}\right).$$

Column 2 shows that the slope is a significant predictor of excess bond returns when considered in isolation. Columns 3 and 4 exploit the decomposition of expected price changes per equation (11) reported here for the reader's convenience:

$$E_t\left(y_{t+4}^{(40-4)}\right) - y_t^{(40-4)} = E_t\left(y_{t+4}^{(40-4),*} - y_t^{(40-4),*}\right) + \left(\rho_{(40-4)} - 1\right) u_t^{(40-4)}.$$

In columns 3 and 4 we neglect the roll-down term which empirically is found to be insignificant. Values in parentheses are conservative standard errors from reverse regressions computed as in Hodrick (1992). Constant estimates are not tabulated. Quarterly observations. The sample period is 1980:Q1 to 2019:Q4. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

	$rx_{t+4}^{(40)}$			
	1	2	3	4
$y_t^{(40)} - y_t^{(4)}$	3.217	2.320*	0.740	0.929
	(2.085)	(1.397)	(2.280)	(1.212)
$-(40-4)\left(E_t\left(y_{t+4}^{(36)}\right) - y_t^{(36)}\right)$	0.543***			
	(0.105)			
$-(40-4)\left(y_t^{(36)} - y_t^{(40)}\right)$	−4.340			
	(4.817)			
$-(40-4)\left(E_t\left(y_{t+4}^{(36),*}\right) - y_t^{(36),*}\right)$			−0.241	
			(2.260)	
$-(40-4)\,u_t^{(36)}$			−0.640***	−0.626***
			(0.164)	(0.119)
Adj. R^2	0.342	0.060	0.316	0.320

predictive regressions when we use bonds with maturities ranging from 2 to 7 years. Also, Table A.3 of the Supplementary Material confirms that stationary deviations of bond prices from their drift predict quarterly holding period bond returns (i.e., nonoverlapping returns). Overall, this evidence suggests that the adjustment of bond prices toward their drift is a key economic mechanism for understanding bond returns predictability.

Finally, Table B.1 of the Supplementary Material shows that the U.S. cyclical component $u_t^{(n)}$ predict U.K. and Canadian bond returns, even after controlling for the local slope of the term structure.[23] This evidence of predictability of international bond risk premia using the U.S. cyclical component is in line with recent work by Miranda-Agrippino and Rey (2020) who find that U.S. monetary policy shocks induce comovements in a single global factor that explains significant variation in financial asset returns.

Our finding also resonates with the evidence in Dahlquist and Hasseltoft (2013). Despite this similarity, Dahlquist and Hasseltoft (2013) attribute the international comovement in bond returns to a global (admittedly, mostly U.S.) bond risk premium; on the other hand, we have not imposed no-arbitrage restrictions so that our cyclical component is also compatible with investors overreacting to

[23]Consistent with our model, we employ the local slope of the term structure as a proxy for the deviations of non-U.S. yields from their drifts. Controlling for the local cyclical component does not change our conclusion. However, we note that the lack of an exogenous potential output series, Δx_t^{pot}, and of a perceived target inflation rate, π_t^*, may be responsible for the poor performance of the local cycle in Canada and U.K. Further investigation on this topic is on our agenda for future research.

https://doi.org/10.1017/S0022109022001557 Published online by Cambridge University Press

deviations of policy rates from its trend leading to overestimation of future short rates (and lower bond returns).[24]

2. The Information Content of Yield Cycles

Several bond returns predictors have been proposed in the literature since the seminal papers by Fama and Bliss (1987) and Campbell and Shiller (1991). It is then natural to ask to what extent the yield cycles $u_t^{(n)}$ capture new information not already conveyed by other variables.

Specifically, we compare the predictive power of our yield cycles to two well-known return-predicting factors that are both constructed from the yield curve[25]: i) the Cochrane and Piazzesi (CP) (2005) factor, which is based on a linear combination of forward rates; and ii) the Cieslak and Povala (CPo) (2015) factor, which relies on information contained in yields that have been detrended using a long-term moving average of inflation.[26] Finally, we also benchmark our yield cycles to the Ludvigson and Ng (2009) single macro-factor – the fitted value from a regression of average (across maturity) excess returns on a set of 6 factors extracted from a large data set of macro-financial variables – which the authors show to contain information about excess bond returns that is not captured by the principal components of the yield covariance matrix.

Rather than using a specific cycle for each maturity n, we construct a common yield cycle using a procedure akin to Cochrane and Piazzesi (2005). Specifically, we run regressions of the average (across maturity) excess return on all cycles,

$$\frac{1}{9}\sum_{n=2}^{10} rx_{t+4}^{(n)} = \gamma_0 + \gamma_1 u_t^{(1)} + \ldots + \gamma_{40} u_t^{(40)} + \varepsilon_{t+1}.$$

Our yield-based cycle factor is given by $\tilde{u}_t = \widehat{\gamma}' \mathbf{u}_t$.

Table 4 shows the results. In Panel A, we investigate the predictive content of our cycle relative to the CP factor, whereas in Panel B we compare it to the CPo factor. The odd columns confirm that both CP and CPo forecasts excess returns of all bonds. Importantly, Panel A shows that our yield cycle drives away the CP factor, and delivers R^2 that are about 3 times those obtained by the CP

[24] Our findings are also consistent with the idea that the Fed is the leader among central banks in setting monetary policy (Brusa, Savor, and Wilson (2019)). See also "One Policy to Rule Them All: Why Central Bank Divergence Is So Slow" (*Wall Street Journal*, Aug. 22, 2016) for a recent discussion on the topic.

[25] Several papers have found that the state of the economy also conveys information about future bond returns. For example, Cooper and Priestley (2008) propose the output gap, whereas Ludvigson and Ng (2009) propose to extract information from a large set of macrofinancial variables. Related, Bansal and Shaliastovich (2013) document that real growth and inflation uncertainties predict, respectively, lower and higher bond risk premia, and propose a long-run risk type model for rationalizing this finding. Since our yield cycles are obtained by removing the stochastic trend (due to the equilibrium rate) in interest rates, we restrict our attention only to yield-based predicting factors.

[26] To construct the CP and CPo factors we follow the procedure described in the original papers. That is, to construct the CP factors we use only 1- through 5-year zero coupon bond prices and estimate the loadings by running a regression of the equal weighted average (across maturity) excess return on the forward rates. To construct the CPo factor instead we employ duration standardized returns. To be consistent with the overall empirical analysis, unlike in the original papers, both factors are constructed using quarterly observations.

TABLE 4

Predictive Regressions: Horse Race Against Other Bond Predictors

Table 4 reports OLS estimates for the regression $rx_{t+4}^{(n)} = \alpha + \beta_1 F_t + \beta_2 \tilde{u}_t + \varepsilon_t$, where $rx_{t+4}^{(n)}$ is the realized 1-year holding period excess return of a bond with maturity n-period, F_t is the Cochrane and Piazzesi (2005) factor (CP_t) in Panel A, the Cieslak and Povala (2015) factor (CPo_t) in Panel B, and the Ludvigson and Ng (2009) factor (LN_t) in Panel C, and $\tilde{u}_t$ is the single-return forecasting factor implied by our model with drifting equilibrium rates. CP_t is constructed as in Cochrane and Piazzesi (2005) using quarterly zero-coupon Treasury yields from Gürkaynak, Sack, and Wright (2007) with maturities from 1 to 5 years. CPo_t is constructed as in Cieslak and Povala (2015) using quarterly zero-coupon Treasury yields from Gürkaynak et al. (2007) with maturities from 1 to 10 years. LN_t is constructed as in Ludvigson and Ng (2009) as the fitted value from a regression of average (across maturity) excess returns on a set of 6 factors extracted from a large data set of macro-financial variables. $\tilde{u}_t$ is the fitted value from regressing the average 1-year holding-period excess returns on a n-periods Treasury bond for $n = 4, 8, \ldots, 40$ on our cyclical components $u_t^{(n)}$, $n = 1, \ldots, 40$ (see equation (11)). Values in parentheses are standard errors from reverse regressions computed as in Hodrick (1992). Constant estimates are not tabulated. Quarterly observations. The sample period is 1980:Q1 to 2019:Q4. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

	$rx_{t+4}^{(8)}$		$rx_{t+4}^{(12)}$		$rx_{t+4}^{(20)}$		$rx_{t+4}^{(28)}$		$rx_{t+4}^{(40)}$	
	1	2	3	4	5	6	7	8	9	10
Panel A. Cochrane–Piazzesi										
CP_t	0.434***	0.104	0.821***	0.168	1.554***	0.340	2.272***	0.589	3.307***	1.069
	(0.139)	(0.148)	(0.305)	(0.331)	(0.580)	(0.633)	(0.810)	(0.887)	(1.141)	(1.263)
$\tilde{u}_t$		0.219***		0.434***		0.807***		1.118***		1.488***
		(0.035)		(0.079)		(0.164)		(0.240)		(0.344)
No. of obs.	156	156	156	156	156	156	156	156	156	156
Adj. R^2	0.130	0.390	0.132	0.421	0.145	0.451	0.159	0.459	0.174	0.446
Panel B. Cieslak–Povala										
CPo_t	1.366***	0.428	2.720***	0.959	5.230***	2.177	7.544***	3.535	10.650***	5.700*
	(0.247)	(0.366)	(0.516)	(0.800)	(1.011)	(1.572)	(1.467)	(2.259)	(2.119)	(3.224)
$\tilde{u}_t$		0.184***		0.346***		0.599**		0.787**		0.972**
		(0.050)		(0.118)		(0.240)		(0.344)		(0.478)
No. of obs.	156	156	156	156	156	156	156	156	156	156
Adj. R^2	0.305	0.396	0.342	0.433	0.389	0.472	0.413	0.487	0.424	0.480
Panel C. Ludvigson and Ng (2009)										
LN_t	0.458	0.365***	0.870***	0.687***	1.472***	1.128**	1.905**	1.419*	2.390**	1.722
	(0.120)	(0.116)	(0.266)	(0.259)	(0.528)	(0.513)	(0.765)	(0.745)	(1.105)	(1.081)
$\tilde{u}_t$		0.221***		0.433***		0.815***		1.153***		1.583***
		(0.034)		(0.074)		(0.151)		(0.218)		(0.310)
No. of obs.	156	156	156	156	156	156	156	156	156	156
Adj. R^2	0.133	0.472	0.136	0.504	0.118	0.518	0.100	0.509	0.080	0.476

regressions. Panel B tells a similar story. Despite the large R^2 obtained by the CPo factor, our yield cycle continues to be a significant predictor of bond returns at all maturities ranging from 2 to 10 years. In fact, comparing the R^2 from the multiple regression in Panel A to those in Panel B, we see that replacing CP with CPo does not alter the predictive content of our yield-based cycle. Finally, in Panel C, we observe that our yield cycle has sizable forecasting power for future excess returns on U.S. government bonds, above and beyond the predictive power contained in the Ludvigson and Ng (2009) single macro-factor.

3. Out-of-Sample Predictability

As a final test, we report the out-of-sample R^2, R^2_{OOS}, computed as follows[27]:

[27]The cointegrating parameters relating the short rate to demographics, productivity, and inflation, are set equal to their values estimated in the full sample (cf Table 1). Thus, the test is pseudo out-of-sample and it is informative of how the model would perform going forward if a practitioner used the existing estimates of these parameters and faced the same distribution of data (see Lettau and Ludvigson (2001)).

TABLE 5
Out-of-Sample Tests

Table 5 reports R^2_{OOS} for the predictive regression $rx^{(n)}_{t+4} = \alpha + \beta' \tilde{u}_t + \varepsilon_t$, where $rx^{(n)}_{t+4}$ is the realized 1-year holding period excess return of a bond with maturity n-period and $\tilde{u}_t$ is the single-return forecasting factor implied by our model with drifting equilibrium rates. $\tilde{u}_t$ is the fitted value from regressing the average 1-year holding-period excess returns on a n-periods Treasury bond for $n = 4, 8, \ldots, 40$ on our cyclical components $u^{(n)}_t$ $n = 1, \ldots, 40$ (see equation (11)). We use a rolling window for estimating the predictive regressions. The R^2_{OOS} is computed as in Campbell and Thompson (2008); p-values for R^2_{OOS} are computed as in Clark and West (2007). In Panel A, the out-of-sample period starts in 1990; in Panel B, the out-of-sample period starts in 2000. Quarterly observations. The sample period is 1980:Q1 to 2019:Q4. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

	$rx^{(8)}_{t+4}$	$rx^{(12)}_{t+4}$	$rx^{(20)}_{t+4}$	$rx^{(28)}_{t+4}$	$rx^{(40)}_{t+4}$
	1	2	3	4	5
Panel A. Out-of-Sample Period: 1990–2019					
R^2_{OOS}	20.9***	27.18***	31.63***	31.65***	27.29***
Panel B. Out-of-Sample Period: 2000–2019					
R^2_{OOS}	0.99***	3.38***	10.23***	14.26***	13.65***

$$R^2_{\text{OOS}} = 1 - \frac{\sum_{t=1}^{T} \left(rx^{(n)}_{t+4} - \widehat{rx}^{(n)}_{t+4} \right)^2}{\sum_{t=1}^{T} \left(rx^{(n)}_{t+4} - \overline{rx}^{(n)}_{t+4} \right)^2},$$

where $\widehat{rx}^{(n)}_{t+4}$ is the fitted value from our predictive regression estimated through period $t-1$ and $\overline{rx}^{(n)}_{t+4}$ is the historical average return estimated through period $t-1$. If the R^2_{OOS} is positive, then the predictive regression has a lower average mean squared prediction error than the historical average return. This is always the case for all regressions reported in Table 5.[28]

IV. Further Results and Discussion

A. The Cyclical Yields Component: A Rational Expectations Interpretation

Our analysis contributes to the debate of whether the bond risk premium is stationary. Hall, Anderson, and Granger (1992) provide early evidence that bond yields are cointegrated and that the bond risk premium is stationary. That view has been challenged, however. For example, Wright (2011) argues for term premiums to decline internationally over the sample 1990–2007. Bauer, Rudebusch, and Wu (2014) and Wright (2014) discuss the extent to which small-sample bias in maximum likelihood estimates of affine term structure models alters the conclusions about term premia and its (a)cyclical properties.

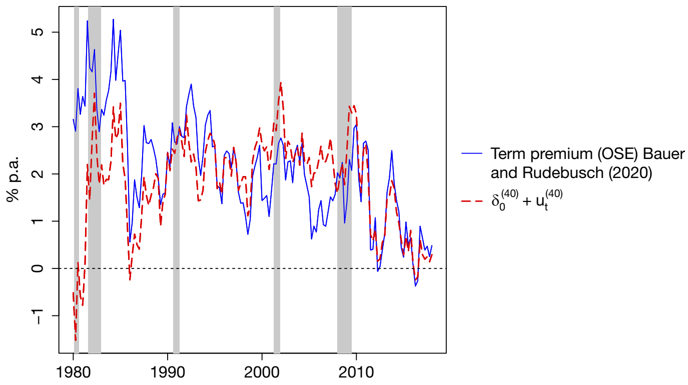
Figure 4 shows the decomposition of the 10-year yield $y^{(40)}_t$ into $y^{(40),*}_t$ and $\delta^{(40)}_0 + u^{(40)}_t$, as per equation (6). As before, Graph A refers to the restricted model whereas Graph B refers to our benchmark model with drifting equilibrium policy rates. The two models have opposite implications: The residuals (dotted line) follow

[28]Panel B of Figure A.4 in the Supplementary Material displays the realized and fitted values for bonds with 2- and 10-years maturity. Also, replacing our proposed drivers for the equilibrium rate (MY, Δx^{pot}_t, and π^*_t) with a time trend would result in negative out-of-sample R^2. See Table A.4 in the Supplementary Material.

FIGURE 5

Cyclical Component from Model with Drifting Equilibrium Rates Versus Term Premium Estimate

Figure 5 shows the term premium component for a 10-year Treasury bond estimated following the methodology (OSE, observed shifting endpoint) proposed by Bauer and Rudebusch (2020) together with deviations of the 10-year bond yields from their drift, $\delta_0^{(40)}+u_t^{(40)}$, implied by our (cointegrated) model with drifting equilibrium rates (cf equations (5)–(8)). Quarterly observations. The sample period is 1980:Q1 to 2018:Q1.



a random walk under the classical model with constant equilibrium rates, but are stationary in our model with drifting rates.[29]

Thus, if we interpret the deviations of bond prices from their drifts as risk premia, we find overwhelming evidence that term premia are indeed stationary. Our evidence complements the literature cited above. In fact, we do not focus on statistical biases but, rather, we stress the importance of modeling the economic determinants of equilibrium rates.

Finally, Figure 5 shows that our estimated deviations of bond prices from their drifts comove strongly with the term premium estimates proposed by Bauer and Rudebusch (2020). This analysis is reminiscent of Joslin, Le, and Singleton (2013) who find that the estimated joint distribution within a macro-finance term structure model with NA is nearly identical to the estimate from an economic-model-free factor vector-autoregression. The evidence in Figure 5 suggests that this conclusion is likely to hold true also in models that accommodate a drifting term structure.

B. The Cyclical Yields Component: A DE Interpretation

In this section, we assess the role played by deviations from REs (in the form of diagnostic expectations) in explaining the cyclical component of yields within our framework.

Under RE, the cyclical component $u_t^{(n)}$ would be identified with the term premium of the n-period bond. However, a stationary $u_t^{(n)}$ is also consistent with, for example, temporary deviations from REs generated within a DE framework

[29]Replacing, in the restricted model, the perceived target rate π_t^* with a fixed target rate at 2%, leaves our conclusion unchanged: the 10-year residual is close to a random walk with an AR(1) coefficient of 0.98.

(see Gennaioli and Shleifer (2018)) where long rates over-react relative to change in expectations about short rates.

Following Bordalo, Gennaioli, and Shleifer (2018) and d'Arienzo (2020), DEs about a stationary process ω_t can be represented as follows:

(12) $$E^{\mathrm{DE}}[\omega_{t+1}|I_t] = E[\omega_{t+1}|I_t] + \theta(E[\omega_{t+1}|I_t] - E[\omega_{t+1}|I_{t-1}]).$$

We apply this expectation formation mechanism to the stationary deviations of the one-period rate from its stochastic trend

(13) $$\omega_{t+1} = y_{t+1}^{(1)} - y_{t+1}^{*}.$$

So, we have

(14) $$E_t^{\mathrm{DE}}\left[y_{t+1}^{(1)} - y_{t+1}^{*}\right] = E_t\left[y_{t+1}^{(1)} - y_{t+1}^{*}\right] + \theta\left(E_t\left[y_{t+1}^{(1)} - y_{t+1}^{*}\right] - E_{t-1}\left[y_{t+1}^{(1)} - y_{t+1}^{*}\right]\right).$$

DEs, $E^{\mathrm{DE}}\left[y_{t+1}^{(1)} - y_{t+1}^{*}|I_t\right]$, differ from REs, $E\left[y_{t+1}^{(1)} - y_{t+1}^{*}|I_t\right]$, by a shift in the direction of the information received at time t on deviations of monetary policy from its (stochastic) trend. Under the DE hypothesis agents overreact to the stationary deviations of monetary policy from its trend.

Since we take the trend in monetary policy rates as exogenous, DEs on the drift coincide with the rational ones. We then write

(15) $$E_t^{\mathrm{DE}}\left[y_{t+1}^{(1)}\right] = E_t\left[y_{t+1}^{(1)}\right] + \theta\left(E_t\left[y_{t+1}^{(1)} - y_{t+1}^{*}\right] - E_{t-1}\left[y_{t+1}^{(1)} - y_{t+1}^{*}\right]\right).$$

Interestingly, in this case, equation (4) can then be rewritten as follows:

(16) $$y_t^{(n)} = \left(\frac{1}{n}\right)\sum_{i=0}^{n-1} E_t\left[y_{t+i}^{(1)}\right] + \delta_0^{(n)} + \underbrace{\left(\frac{1}{n}\right)\sum_{i=0}^{n-1}\left(E_t^{\mathrm{DE}}\left[y_{t+i}^{(1)}\right] - E_t\left[y_{t+i}^{(1)}\right]\right)}_{u_t^{(n)}}.$$

Thus, the (stationary) component $u_t^{(n)}$ can in principle be explained by the over-reaction induced by DE: that is, $u_t^{(n)}$ can be justified also if term premia are constant or even absent.

We rewrite equation (16) as follows:

(17) $$\begin{aligned} y_t^{(n)} &= \left(\frac{1}{n}\right)\sum_{i=0}^{n-1} E_t\left[y_{t+i}^{(1)}\right] + \delta_0^{(n)} + \left(\frac{1}{n}\right)\sum_{i=0}^{n-1}\left(E_t^{\mathrm{DE}}\left[y_{t+i}^{(1)}\right] - E_t\left[y_{t+i}^{(1)}\right]\right) \\ &= \left(\frac{1}{n}\right)\sum_{i=0}^{n-1} E_t\left[y_{t+i}^{(1)}\right] + \delta_0^{(n)} + \\ &+ \left(\frac{1}{n}\right)\theta\sum_{i=0}^{n-1}\left(E_t\left[y_{t+i}^{(1)} - y_{t+i}^{*}\right] - E_{t-1}\left[y_{t+i}^{(1)} - y_{t+i}^{*}\right]\right), \end{aligned}$$

where in the second step we exploit equation (15).

TABLE 6

Testing Diagnostic Expectations

Table 6 reports seemingly unrelated regressions (SUR) estimates for regression (equation (17)) for bonds with different maturities (n). We restrict θ to be the same across maturities. Quarterly observations. The sample period is 1980:Q1 to 2019:Q4. ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

	$u_t^{(8)}$	$u_t^{(12)}$	$u_t^{(20)}$	$u_t^{(28)}$	$u_t^{(40)}$
	1	2	3	4	5
$1/n\sum_{i=0}^{n-4}(E_t - E_{t-1})\left[y_{t+i}^{(1)} - y_{t+i}^{*}\right]$	3.679***	3.679***	3.679***	3.679***	3.679***
	(0.299)	(0.299)	(0.299)	(0.299)	(0.299)
Constant	0.717***	0.877***	1.189***	1.467***	1.798***
	(0.058)	(0.067)	(0.076)	(0.081)	(0.085)
No. of obs.	160	160	160	160	160
Adj. R^2	0.408	0.254	0.112	0.057	0.026

In all, we can estimate the θ parameter by running the following regression:

$$y_t^{(n)} - y_t^{(n,*)} = \delta_0^{(n)} + \theta\left(\frac{1}{n}\right)\sum_{i=0}^{n-1}(E_t - E_{t-1})\left[y_{t+i}^{(1)} - y_{t+i}^{*}\right],$$

where $\left(\frac{1}{n}\right)\theta\sum_{i=0}^{n-1}\left(E_t\left[y_{t+i}^{(1)} - y_{t+i}^{*}\right] - E_{t-1}\left[y_{t+i}^{(1)} - y_{t+i}^{*}\right]\right)$ is obtained from forward simulation of our model.

Several comments are in order. First, deviations from REs depend on the parameter θ and on the persistence of the deviations of monetary policy rates from the trend. Second, the stationarity of $\left(y_t^{(1)} - y_t^{*}\right)$ implies that, for large n (i.e., at long horizons), DEs for the monetary policy rates will converge toward REs.[30] Third, and most important, the estimated value of the θ parameter allows to assess the relevance of DEs under the null of our model.

Table 6 displays the results for such test. We find that DEs explain between 3% and 40% of the variability in the cyclical components of yields for bonds with maturity from 2 to 10 years. In line with our discussion of equation (17), the importance of DEs decreases with the maturity of the bond.

The empirical relevance of overreaction has been recently documented by Cieslak (2018) for the short end of the curve. Similarly, Piazzesi, Salomao, and Schneider (2015) provide evidence that realized survey (interest rates) forecast errors as well as forecast differences relative to VAR-based measure may be responsible for the time-variation in bond premia from statistical models. We have shown that these explanations may be important even in a model that accommodates a drifting term structure. However, the contribution of overeaction decreases at long maturities; this is consistent with deviations of monetary policy rates from the equilibrium rate being fast mean-reverting.

[30]Maxted (2019) considers a case in which convergence of DE to RE is not realized as the underlying process is nonstationary.

V. Conclusion

This article proposes a general framework to model a common drift in bond prices, and studies its implications for monetary policy, the term structure of interest rates, and bond returns predictability.

We start by showing that there is a drift in monetary policy rates which can be successfully modeled by fluctuations in productivity, demographics, and long-term inflation expectations. Our approach delivers monetary policy residuals that are substantially less persistent than those implied by standard policy rules. Thus, through the lens of our analysis, we find that monetary policy inertia is over-estimated when the drift in policy rate is not modeled.

The drift in bond prices is described by the average of expected monetary policy rates over the residual life of the bond. Appropriate modeling of the drift in monetary policy should deliver stationary deviation of yields to maturity from their drift. We find that persistent but stationary deviations of U.S. Treasury bond prices from their drift predict excess returns in- and out-of-sample, as well as outside the United States. Next period returns from holding long-term bonds are negative in times when bond prices are higher than those implied by their drift.

Finally, stationary deviations of bond prices from their drift could be explained by the presence of term premia and/or by temporary deviations from REs in a behavioral framework. Our empirical evidence shows that deviations from REs in the form of DEs account for up to 40% of the fluctuations in yield cycles for bonds with maturities 2-year. However, the importance of DEs decreases at longer maturities leaving an important role for term premia. At a minimum, when the deviations of bond prices from their drift are interpreted as term premia, our finding implies that models that misspecify the drift in monetary policy and in bond prices will fail to generate stationary term premia.

Appendix. Data

We employ quarterly data in our empirical analysis; thus, we proxy for the 1-period bond yields using the end-of-quarter 3-month Treasury bill rates from the Federal Reserve's H.15 release. Our sample period starts with Paul Volcker's appointment as Fed chairman, because of evidence that monetary and macroeconomic dynamics changed at that time (e.g., Gertler, Gali, and Clarida (1999)).

Zero-coupon Treasury yields with 1- to 10-year maturities are from Gürkaynak et al. (2007).

The Federal Reserve's perceived target rate (PTR) for inflation is a survey-based measure of long-run inflation expectations; PTR is used in the Fed's FRB/U.S. model and is available at https://www.federalreserve.gov/econres/us-models-package.htm.

MY is available until 2050 and is hand-collected from various past Census reports available at https://www.census.gov/data.html. Potential output is available until 2030 and can be downloaded at https://fred.stlouisfed.org/series/GDPPOT. See also Figure A1.

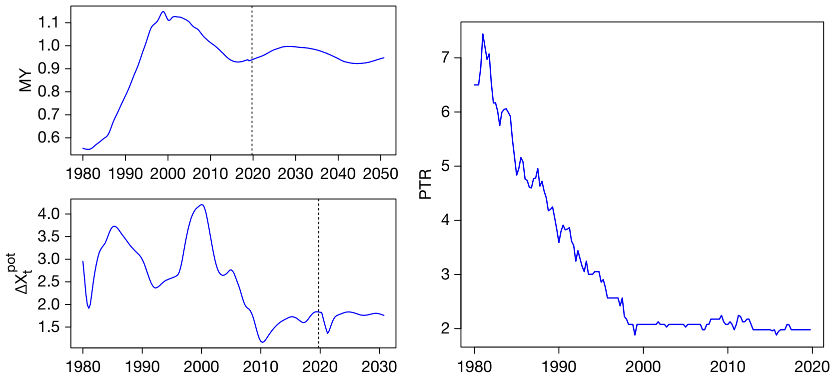
A natural concern for our analysis of bond return predictability is that Δx^{pot} may be formed in a way that exploits information from the yield curve.[31] Fortunately, this is not

[31]We thank the referee for raising this point.

FIGURE A1

Drivers of the Equilibrium Nominal Rate

Figure A1 shows the dynamics for the drivers of the time-varying equilibrium nominal rate y_t^* in equation (2). The left graph shows the ratio of middle-aged (40–49) to young (20–29) population, MY, and for potential output growth, Δx_t^{pot}. The right graph shows the Federal Reserve's perceived target rate (PTR) for inflation. MY is available until 2050 and is hand-collected from various past Census reports available at https://www.census.gov/data.html. Potential output is available until 2030 and can be downloaded at https://fred.stlouisfed.org/series/GDPPOT. Dotted vertical lines denote the end of our sample, that is, 2019:Q4. Quarterly observations.



the case. Indeed, the U.S. Congressional Budget Office (CBO) defines potential output as the trend growth in the productive capacity of the economy. CBO uses the Solow growth model in which the real GDP growth is the product of 3 input factors: capital, labor, and technology. They attribute GDP to 5 sectors: nonfarm business, government, farm, households and nonprofits, and housing. For every sector, the CBO estimates a standard production function based on labor, capital, and total factor productivity (TFP). Aggregate real GDP is the sum of real GDP across the 5 sectors.

To estimate potential output, CBO estimates for each sector the potential level of labor force $N_{i,t}^*$ which is a function of the unemployment gap and business cycle dummies. With $N_{i,t}^*$ in hands, CBO computes the potential level of hours worked $L_{i,t}^*$. Then, CBO cyclically adjusts TFP to remove business cycle fluctuations. Finally, they compute for each sector potential output as $Y_{i,t}^* = A_{i,t} L_{i,t}^{*(0.7)} K_{i,t}^{(0.3)}$. For a complete description of the methodology used by the CBO, see https://www.cbo.gov/sites/default/files/107th-congress-2001-2002/reports/potentialoutput.pdf. In all, the CBO does not look at the yield curve when forming Δx_t^{pot}.

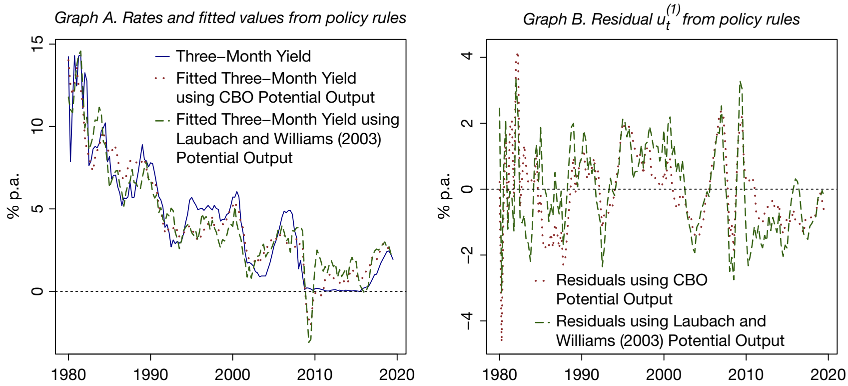
We also replace Δx^{pot} from the CBO with the potential output provided by Laubach and Williams (2003). Figure A2 shows the results. In particular, Graph A shows the 23-month yield and fitted values for our (cointegrated) model that uses either the CBO or the Laubach and Williams (2003) potential output. Graph B shows the cycle $u_t^{(1)}$.

Figure A2 shows that the 2 fitted values for the short-term rate (Graph A) and the implied cycles (Graph B) are highly correlated. This evidence suggests that the adoption of the potential output from CBO is not essential to obtain the results.

FIGURE A2

Actual Versus Fitted Short-Term Rate and Cycle: CBO Versus Laubach and Williams (2003) Potential Output

Graph A of Figure A2 shows the actual 23-month yield and fitted values for our (cointegrated) model with drifting equilibrium rates (cf equation (2)). The red dotted line is the benchmark case in the paper based on CBO potential output. The green dashed line is instead the fitted value from a model that employs the potential output provided by Laubach and Williams (2003). Graph B shows the differences between the actual 23-months yield and the fitted values. Quarterly observations. The sample period is 1980:Q1 to 2019:Q4.



Supplementary Material

To view supplementary material for this article, please visit http://doi.org/10.1017/S0022109022001557.

References

Ang, A.; J. Boivin; S. Dong; and R. Loo-Kung. "Monetary Policy Shifts and the Term Structure." *Review of Economic Studies*, 78 (2011), 429–457.

Aruoba, S. B., and F. Schorfheide. "Inflation During and After the Zero Lower Bound." In *2015 Jackson Hole Symposium Volume*, Federal Reserve Bank of Kansas City (2016), 359–436.

Balduzzi, P.; S. R. Das; and S. Foresi. "The Central Tendency: A Second Factor in Bond Yields." *Review of Economics and Statistics*, 80 (1998), 62–72.

Bansal, R., and I. Shaliastovich. "A Long-Run Risks Explanation of Predictability Puzzles in Bond and Currency Markets." *Review of Financial Studies*, 26 (2013), 1–33.

Bauer, M. D., and G. D. Rudebusch. "Interest Rates Under Falling Stars." *American Economic Review*, 110 (2020), 1316–1354.

Bauer, M. D.; G. D. Rudebusch; and J. C. Wu. "Term Premia and Inflation Uncertainty: Empirical Evidence from an International Panel Dataset: Comment." *American Economic Review*, 104 (2014), 323–337.

Berardi, A.; M. Markovich; A. Plazzi; and A. Tamoni. "Mind the (Convergence) Gap: Bond Predictability Strikes Back!" *Management Science*, 67 (2020), 7291–7950.

Bernanke, B. S., and K. N. Kuttner. "What Explains the Stock Market's Reaction to Federal Reserve Policy?" *Journal of Finance*, 60 (2005), 1221–1257.

Bordalo, P.; N. Gennaioli; and A. Shleifer. "Diagnostic Expectations and Credit Cycles." *Journal of Finance*, 73 (2018), 199–227.

Brusa, F.; P. Savor; and M. Wilson. "One Central Bank to Rule Them All." *Review of Finance*, 24 (2019), 263–304.

Campbell, J. "Long-Term Investing in a Nonstationary World." NBER Lecture (2019).

Campbell, J. Y., and R. J. Shiller. "The Dividend-Price Ratio and Expectations of Future Dividends and Discount Factors." *Review of Financial Studies*, 1 (1988), 195–228.

Campbell, J. Y., and R. J. Shiller. "Yield Spreads and Interest Rate Movements: A Bird's Eye View." *Review of Economic Studies*, 58 (1991), 495–514.

Campbell, J. Y., and R. Sigalov. "Portfolio Choice with Sustainable Spending: A Model of Reaching for Yield." NBER Technical Report (2020).
Campbell, J. Y., and S. B. Thompson. "Predicting Excess Stock Returns Out of Sample: Can Anything Beat the Historical Average?" *Review of Financial Studies*, 21 (2008), 1509–1531.
Chava, S., and A. Hsu. "Financial Constraints, Monetary Policy Shocks, and the Cross-Section of Equity Returns." *Review of Financial Studies*, 33 (2020), 4367–4402.
Cieslak, A. "Short-Rate Expectations and Unexpected Returns in Treasury Bonds." *Review of Financial Studies*, 31 (2018), 3265–3306.
Cieslak, A., and P. Povala. "Expected Returns in Treasury Bonds." *Review of Financial Studies*, 28 (2015), 2859–2901.
Clarida, R.; J. Gali; and M. Gertler. "Monetary Policy Rules and Macroeconomic Stability: Evidence and Some Theory." *Quarterly Journal of Economics*, 115 (2000), 147–180.
Clark, T., and K. West. "Approximately Normal Tests for Equal Predictive Accuracy in Nested Models." *Journal of Econometrics*, 138 (2007), 291–311.
Cochrane, J. H. "Comments on 'Macroeconomic Implications of Changes in the Term Premium'." *Federal Reserve Bank of St. Louis Review*, 89 (2007), 271–282.
Cochrane, J. H., and M. Piazzesi. "Bond Risk Premia." *American Economic Review*, 95 (2005), 138–160.
Cogley, T.; G. E. Primiceri; and T. J. Sargent. "Inflation-Gap Persistence in the US." *American Economic Journal: Macroeconomics*, 2 (2010), 43–69.
Coibion, O., and Y. Gorodnichenko. "Monetary Policy, Trend Inflation, and the Great Moderation: An Alternative Interpretation." *American Economic Review*, 101 (2011), 341–370.
Coibion, O., and Y. Gorodnichenko. "Why Are Target Interest Rate Changes So Persistent?" *American Economic Journal: Macroeconomics*, 4 (2012), 126–162.
Cooper, I., and R. Priestley. "Time-Varying Risk Premiums and the Output Gap." *Review of Financial Studies*, 22 (2008), 2801–2833.
d'Arienzo, D. "Increasing Overreaction and Excess Volatility of Long Rates." Working Paper, Bocconi University (2020).
Dahlquist, M., and H. Hasseltoft. "International Bond Risk Premia." *Journal of International Economics*, 90 (2013), 17–32.
Duffee, G. R. "Term Premia and Interest Rate Forecasts in Affine Models." *Journal of Finance*, 57 (2002), 405–443.
Duffee, G. R. "Forecasting Interest Rates." In *Handbook of Economic Forecasting, Vol. 2 of Handbook of Economic Forecasting*, G. Elliott, C. Granger, and A. Timmermann, eds. Amsterdam, Netherlands: Elsevier (2013), 385–426.
Engle, R. F., and C. W. Granger. "Co-Integration and Error Correction: Representation, Estimation, and Testing." *Econometrica*, 55 (1987), 251–276.
Fama, E. F. "The Behavior of Interest Rates." *Review of Financial Studies*, 19 (2006), 359–379.
Fama, E. F., and R. R. Bliss. "The Information in Long-Maturity Forward Rates." *American Economic Review*, 77 (1987), 680–692.
Farhi, E., and F. Gourio. "Accounting for Macro-Finance Trends: Market Power, Intangibles, and Risk Premia." NBER Technical Report (2018).
Feunou, B., and J.-S. Fontaine. "Secular Economic Changes and Bond Yields." *Review of Economics and Statistics*, forthcoming, https://doi.org/10.1162/rest_a_01034 (2023).
Geanakoplos, J.; M. Magill; and M. Quinzii. "Demography and the Long-Run Predictability of the Stock Market." *Brookings Papers on Economic Activity*, 2004 (2002), 241–325.
Gennaioli, N., and A. Shleifer. *A Crisis of Beliefs: Investor Psychology and Financial Fragility*. Princeton, NJ: Princeton University Press (2018).
Gertler, M.; J. Gali; and R. Clarida. "The Science of Monetary Policy: A New Keynesian Perspective." *Journal of Economic Literature*, 37 (1999), 1661–1707.
Giammarioli, N., and N. Valla. "The Natural Real Interest Rate and Monetary Policy: A Review." *Journal of Policy Modeling*, 26 (2004), 641–660.
Giannone, D.; M. Lenza; and G. E. Primiceri. "Priors for the Long Run." *Journal of the American Statistical Association*, 114 (2019), 565–580.
Golinski, A., and P. Zaffaroni. "Long Memory Affine Term Structure Models." *Journal of Econometrics*, 191 (2016), 33–56.
Gürkaynak, R. S.; B. Sack; and J. H. Wright. "The US Treasury Yield Curve: 1961 to the Present." *Journal of Monetary Economics*, 54 (2007), 2291–2304.
Hall, A. D.; H. M. Anderson; and C. W. J. Granger. "A Cointegration Analysis of Treasury Bill Yields." *Review of Economics and Statistics*, 74 (1992), 116–126.
Hodrick, R. J. "Dividend Yields and Expected Stock Returns: Alternative Procedures for Inference and Measurement." *Review of Financial Studies*, 5 (1992), 357–386.

Jardet, C.; A. Monfort; and F. Pegoraro. "No-Arbitrage Near-Cointegrated VAR(p) Term Structure Models, Term Premia and GDP Growth." *Journal of Banking & Finance*, 37 (2013), 389–402.
Jordà, Ò., and A. M. Taylor. "Riders on the Storm." NBER Technical Report (2019).
Jørgensen, K. "How Learning from Macroeconomic Experiences Shapes the Yield Curve." Finance and Economics Discussion Series 2018, Board of Governors of the Federal Reserve System (2018).
Joslin, S.; A. Le; and K. J. Singleton. "Why Gaussian Macro-Finance Term Structure Models are (nearly) Unconstrained Factor-VARs." *Journal of Financial Economics*, 109 (2013), 604–622.
Kiley, M. T. "What Can the Data Tell Us About the Equilibrium Real Interest Rate?" Finance and Economics Discussion Series 2015-77, Board of Governors of the U. S. Federal Reserve System (2015).
Koijen, R. S.; H. Lustig; and S. Van Nieuwerburgh. "The Cross-Section and Time Series of Stock and Bond Returns." *Journal of Monetary Economics*, 88 (2017), 50–69.
Kozicki, S., and P. A. Tinsley. "Shifting Endpoints in the Term Structure of Interest Rates." *Journal of Monetary Economics*, 47 (2001), 613–652.
Laubach, T., and J. C. Williams. "Measuring the Natural Rate of Interest." *Review of Economics and Statistics*, 85 (2003), 1063–1070.
Lettau, M., and S. Ludvigson. "Consumption, Aggregate Wealth, and Expected Stock Returns." *Journal of Finance*, 56 (2001), 815–849.
Ludvigson, S. C., and S. Ng. "Macro Factors in Bond Risk Premia." *Review of Financial Studies*, 22 (2009), 5027–5067.
Lunsford, K. G., and K. D. West. "Some Evidence on Secular Drivers of US Safe Real Rates." *American Economic Journal: Macroeconomics*, 11 (2019), 113–139.
MacKinnon, J. G. "Critical Values for Cointegration Tests." Technical Report, Queen's Economics Department (2010).
Maxted, P. "A Macro-Finance Model with Sentiment." Working Paper, Harvard University (2019).
Miranda-Agrippino, S., and H. Rey. "U.S. Monetary Policy and the Global Financial Cycle." *Review of Economic Studies*, 87 (2020), 2754–2776.
Negro, M. D.; D. Giannone; M. P. Giannoni; and A. Tambalotti. "Safety, Liquidity, and the Natural Rate of Interest." *Brookings Papers on Economic Activity*, 48 (2017), 235–316.
Orphanides, A. "Historical Monetary Policy Analysis and the Taylor Rule." *Journal of Monetary Economics*, 50 (2003), 983–1022.
Ortu, F.; F. Severino; A. Tamoni; and C. Tebaldi. "A Persistence-Based Wold-Type Decomposition for Stationary Time Series." *Quantitative Economics*, 11 (2020), 203–230.
Ozdagli, A. K. "Financial Frictions and the Stock Price Reaction to Monetary Policy." *Review of Financial Studies*, 31 (2018), 3895–3936.
Pasten, E.; R. Schoenle; and M. Weber. "The Propagation of Monetary Policy Shocks in a Heterogeneous Production Economy." *Journal of Monetary Economics*, 116 (2020), 1–22.
Phillips, P. C., and P. Perron. "Testing for a Unit Root in Time Series Regression." *Biometrika*, 75 (1988), 335–346.
Piazzesi, M., and J. Cochrane. "Decomposing the Yield Curve." Technical Report, University of Chicago (2008).
Piazzesi, M.; J. Salomao; and M. Schneider. "Trend and Cycle in Bond Premia." Working Paper, Stanford University (2015).
Rudebusch, G. D. "Term Structure Evidence on Interest Rate Smoothing and Monetary Policy Inertia." *Journal of Monetary Economics*, 49 (2002), 1161–1187.
Rudebusch, G. D. "Monetary Policy Inertia: Fact or Fiction?" Technical Report 4, Federal Reserve Bank of San Francisco (2006).
Sarno, L.; P. Schneider; and C. Wagner. "The Economic Value of Predicting Bond Risk Premia." *Journal of Empirical Finance*, 37 (2016), 247–267.
Swanson, E. T., and J. C. Williams. "Measuring the Effect of the Zero Lower Bound on Medium-and Longer-Term Interest Rates." *American Economic Review*, 104 (2014), 3154–85.
Taylor, J. B., "Discretion Versus Policy Rules in Practice." *Carnegie-Rochester Conference Series on Public Policy*, 39 (1993), 195–214.
van Binsbergen, J. H. "Duration-Based Stock Valuation." NBER Working Paper No. 27367 (2020).
Woodford, M. "The Taylor Rule and Optimal Monetary Policy." *American Economic Review*, 91 (2001), 232–237.
Woodford, M. "Optimal Interest-Rate Smoothing." *Review of Economic Studies*, 70 (2003), 861–886.
Wright, J. H. "Term Premia and Inflation Uncertainty: Empirical Evidence from an International Panel Dataset." *American Economic Review*, 101 (2011), 1514–1534.
Wright, J. H. "Term Premia and Inflation Uncertainty: Empirical Evidence from an International Panel Dataset: Reply." *American Economic Review*, 104 (2014), 338–341.