Journal of Macroeconomics 75 (2023) 103484

Contents lists available at ScienceDirect

Journal of Macroeconomics

journal homepage: www.elsevier.com/locate/jmacro

Yield curve and the macroeconomy: Evidence from a DSGE model with housing[☆]

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ARTICLE INFO

JEL classification:
E23
E32
E44
E52
R31

Keywords:
DSGE model
Yield curve
Housing
Regime switching
Nominal rigidities
Adjustment costs

ABSTRACT

The slope of the yield curve has long been found to be a useful predictor of future economic activities, but the relationship is unstable. One change we have identified in this paper is that, between the early 1990s and the collapse of the housing market in 2007, movements at the long end of the yield curve have an increase in predictive power. We use a medium-scale DSGE model with a housing sector and a yield curve as a guide to find out the sources of such change. The model implies that an increase in the short-term interest rate and a decrease in the long-term interest rate have different impacts on the economy, and to use the slope as a predictor one needs to distinguish movements at the two ends of the yield curve. Based on simulated data from the model, we find that nominal wage rigidities and the capital adjustment costs are closely related to the predictive power of the yield curve. This result is further confirmed with actual data.

1. Introduction

The practice of using interest rate spreads as macroeconomic predictors can be dated back to Stock and Watson (1989) who include the long-term and short-term treasury bond yield spread as one of the leading economic indicators. Subsequent empirical studies, including Estrella and Hardouvelis (1991), Estrella and Mishkin (1996), Dueker (1997), and Estrella and Mishkin (1998), among many others, all find that the spread between 10-year and 3-month interest rates is a useful predictor for future recessions and its predictive power significantly outperforms other financial and macroeconomic indicators.

In recent years, there has been renewed interest in the yield spread as a predictor of future economic activities. Chinn and Kucko (2015) examine the predictive power of the yield spread across a group of advanced economies and over various sample periods. Using a probit regression, they find that the yield spread is a strong predictor of future recessions before the 2000s but the predictive power has deteriorated prior to the Great Recession. Walsh (2010) and Aguirre and Vázquez (2020) also show that the correlation between interest rates and output becomes less pronounced since the mid 1980s. Motivated by Wright (2006)'s argument that there is no fundamental reason why movements at the two ends of the yield curve must have the same predictive content for the likelihood of a recession, Chinn and Kucko (2015) include the short-term policy rate in addition to the yield spread in the regression, so that changes in the yield spread capture movements at the long end of the yield curve.

☆ The authors would like to thank Dr. Marco Del Negro for his valuable comments on an earlier version of this paper. We are also grateful for the insightful comments and suggestions made by the Editor, Dr. David VanHoose, and three anonymous referees.

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https://doi.org/10.1016/j.jmacro.2022.103484
Received 17 December 2021; Received in revised form 28 July 2022; Accepted 14 November 2022
Available online 21 November 2022
0164-0704/

Table 1
Yield spread as predictor of NBER recessions over the next four quarters.

Bai & Perron test	1% critical value			5% critical value			10% critical value		
3773.50	5.13			4.18			3.77		
Break points	Point estimate			95% confidence interval					
Point 1	1990:Q3			1990:Q2			1990:Q4		
Point 2	2000:Q1			1970:Q3			2029:Q3		
Point 3	2007:Q1			2006:Q1			2008:Q1		
	1972:Q1–1990:Q3			1990:Q4–2007:Q1			2007:Q2–2019:Q4		
Spread	−413.19***	−370.06***	−461.27***	−338.18***	−301.22*	−473.64	−227.85***	197.72	−392.69***
	(−6.38)	(−4.51)	(−5.47)	(−2.76)	(−1.85)	(−1.46)	(−2.85)	(0.91)	(−3.26)
Short rate		91.20			172.41			590.40**	
		(1.51)			(0.93)			(2.50)	
Long rate			91.20			172.41			590.40**
			(1.51)			(0.93)			(2.50)
N	75	75	75	67	67	67	51	51	51
Mc Fadden R^2	0.544	0.583	0.583	0.281	0.335	0.335	0.283	0.559	0.559

Entries in parentheses are t statistics constructed from Newey–West standard errors. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

We present similar empirical results in Table 1 based on the following simple probit regressions:

$$\Pr(\mathrm{R}_{t+1,t+h} = 1) = \Phi(\beta_0 + \beta_1 \mathrm{Spread}_t + \epsilon_{t+h}), \quad (1)$$

$$\Pr(\mathrm{R}_{t+1,t+h} = 1) = \Phi(\beta_0 + \beta_1 \mathrm{Spread}_t + \beta_2 \text{Short Rate}_t + \epsilon_{t+h}), \quad (2)$$

$$\Pr(\mathrm{R}_{t+1,t+h} = 1) = \Phi(\beta_0 + \beta_1 \mathrm{Spread}_t + \beta_2 \text{Long Rate}_t + \epsilon_{t+h}), \quad (3)$$

using quarterly data, where the recession indicator variable $R_{t+1,t+h}$ equals one if there is a recession identified by the NBER's Business Cycle Dating Committee in any quarter between $t+1$ and $t+h$, inclusively. We consider a forecast horizon of $h = 4$, over which the yield spread has been found to have important and significant predictive power in general; see Chinn and Kucko (2015). We use the 3-month Treasury Bill secondary market rate obtained from the Board of Governors of the Federal Reserve System as the short-term interest rate. Given that the short-term interest rate stays near zero in the aftermath of the 2008 financial crisis, we add to it the difference between the (Wu and Xia, 2016) shadow federal funds rate and the effective federal funds rate of the Federal Reserve after 2008 to account for the effects of unconventional monetary policy. The long-term interest rate is the 10-year zero-coupon yields calculated by Gürkaynak, Sack, and Wright (2007). The yield spread is the difference between the long- and short-term interest rates. Our sample ranges from 1972:Q1 to 2019:Q4. The start of the sample is restricted by the data availability of the long-term interest rate and the end of the sample is chosen to exclude the economic turbulence caused by the COVID-19 pandemic.

In the top panel of Table 1, we conduct a test for multiple unknown structural breaks based on Eq. (2) using the Bai and Perron (1998, 2003) method. The test statistic well exceeds the critical value at the 1% level and the test suggests three break points at 1990:Q3, 2000:Q1, and 2007:Q1, respectively. However, the second break point is estimated to have an unreasonably wide confidence interval. We therefore split the sample at the two break points, 1990:Q3 and 2007:Q1, and present the subsample estimation results in the bottom panel of the table.

In the early subsample period 1972:Q1–1990:Q3, the yield spread alone is a highly significant predictor of future recessions. A flattening yield curve predicts a higher probability of recessions over the next four quarters. The overall model fit of the simple probit regression in Eq. (1) as measured by the McFadden R^2 exceeds 0.5. The predictability turns out to rely more on movements at the short end than the long end of the yield curve, as the estimated coefficient of the yield spread in Eq. (3) exceeds that in Eq. (2) in magnitude. The predictive power of the yield spread deteriorates over the second subsample period 1990:Q4–2007:Q1 and the R^2 reduces to 0.28. After controlling for the short-term interest rate, however, the model fit improves. The yield spread remains a significant predictor of future recessions in Eq. (2) where the short-term interest rate is controlled but becomes insignificant in Eq. (3) which instead controls for the long-term interest rate. This result suggests that movements at the long end of the yield curve outweigh those at the short end in predicting future economic activities over the second subsample period, which is not surprising given the fact that it covers the housing boom leading up to the financial crisis. The demand side of the housing market is heavily affected by the long-term interest rate, because, as the Monthly Interest Rate Survey of the Federal Housing Finance Agency shows, about 78% of the mortgages over that period of time are fixed-rate loans with an average maturity of 27 years.[1] Over the third subsample period 2007:Q2–2019:Q4, however, movements at the short end of the yield curve regain predictive power.

While all of the aforementioned studies employ reduced-form regressions to examine the predictive power of the yield spread, in this paper we try to understand the predictive power of the yield spread through the lens of a medium-scale structural macroeconomic model. The incorporation of the yield spread into a structural model enables us to study the underlying mechanisms

[1] See Table 17 in https://www.fhfa.gov/DataTools/Downloads/Pages/Monthly-Interest-Rate-Data.aspx.

through which the yield spread might affect future economic activities. Motivated by the empirical evidence that the long end of yield curve outperforms the short end in predicting future recessions over a sample period that includes the drastic rise in housing price, we also include a housing sector in our structural model. To our best knowledge, this is the first paper that studies the yield spread as predictor of future recessions in a structural framework.

Through a series of counterfactual analyses based on our estimated model, we find that both nominal wage rigidities and capital adjustment costs affect the predictive power of the yield spread. With higher wage rigidities or lower capital adjustment costs, both residential and nonresidential investment becomes more responsive to structural shocks, and therefore a decrease in the yield spread predicts a larger increase in the likelihood of future recessions. The occurrence between the early 1990s and the collapse of the housing market in 2007 of an increase in nominal wage rigidities and a decrease in capital adjustment costs is identified using a regime-switching version of the model, which matches our empirical finding that movements at the long end of the yield curve exhibit more predictive power over this sample period.

The remainder of the paper is as follows. Section 2 introduces the model economy. Section 3 takes the model to data. Section 4 examines potential factors impacting the predictive power of yield spread through a series of counterfactual analyses. Section 5 introduces regime switching to key parameters in the model economy. Section 6 provides further evidence from actual data. Section 7 concludes the paper.

2. The model economy

We construct a model economy that features a housing sector based on Iacoviello and Neri (2010). On the demand side, there are two types of households with different discount factors — (credit-unconstrained) lenders and (credit-constrained) borrowers. On the supply side, entrepreneurs produce consumption goods and new houses. Both entrepreneurs and constrained households borrow from unconstrained households subject to collateral constraints, with the difference that entrepreneurs borrow at the short rate and constrained households borrow at the long rate.[2] Our specification of the yield curve follows De Graeve, Emiris, and Wouters (2009) where, under the expectation hypothesis, long-term interest rates reflect economic agents' beliefs about the future path of the short-term rate or policy rate.

2.1. Entrepreneurs

There is a continuum of measure one of entrepreneurs who accumulate capital and own land. They hire labor and purchase intermediate goods to produce wholesale goods Y_t and new houses IH_t. Let c, b, k_c and k_h denote consumption, borrowing, and capital stocks in the consumption and housing sectors. Let a_k capture the investment-specific shock that follows an AR(1) process and represent the marginal cost of producing capital used in the consumption sector. Loans are set in nominal terms and entrepreneurs borrow at a short-term interest rate R_S using land l as the collateral asset. Money inflation in the consumption sector is π. Entrepreneurs hire labor, from unconstrained and constrained households, n'_c and n''_c for the production of consumption goods and n'_h and n''_h for the production of new houses. Variables and parameters with a single prime refer to unconstrained households and those with double primes refer to constrained households. Real wages are at w'_c, w''_c, w'_h, and w''_h. Entrepreneurs also act as a financial intermediary between unconstrained and constrained households and earn the difference between long-term and short-term interest rates $R_L - R_S$. Intermediate input, land prices, and real house prices are denoted k_b, p_l, and q, respectively. Retailers purchase consumption goods from entrepreneurs and then sell them at a markup X. Let z_t capture the shock to inter-temporal preferences which follows an AR(1) process. Entrepreneurs maximize their lifetime utility:

$$V = E_0 \sum_{t=0}^{\infty} (\beta)^t z_t \frac{1-\epsilon}{1-\beta\epsilon} \ln(c_t - \epsilon c_{t-1}), \tag{4}$$

subject to the following budget constraint:

$$\begin{aligned}
&\frac{Y_t}{X_t} + q_t IH_t + b_t + \frac{X_t - 1}{X_t} Y_t + \frac{b''_{t-1}(R_{L,t-1} - R_{S,t-1})}{\pi_t} \\
=&c_t + \frac{R_{S,t-1} b_{t-1}}{\pi_t} + \frac{k_{c,t} - (1-\delta_{kc}) k_{c,t-1}}{a_{k,t}} + \left(k_{h,t} - (1-\delta_{kh}) k_{h,t-1}\right) \\
&+ k_{b,t} + p_{l,t}(l_t - l_{t-1}) + \sum_{i=c,h} w'_{i,t} n'_{i,t} + \sum_{i=c,h} w''_{i,t} n''_{i,t} \\
&+ \frac{\phi_{kc}}{2} \left(\frac{k_{c,t}}{k_{c,t-1}} - 1 \right)^2 k_{c,t-1} + \frac{\phi_{kh}}{2} \left(\frac{k_{h,t}}{k_{h,t-1}} - 1 \right)^2 k_{h,t-1},
\end{aligned} \tag{5}$$

and a collateral constraint:

$$b_t \le m E_t \left(p_{l,t+1} l_t \frac{\pi_{t+1}}{R_{S,t}} \right). \tag{6}$$

[2] The Survey of Business Lending conducted by the Board of Governors of the Federal Reserve System shows that the weighted-average maturity for all commercial and industry loans is between 8 months and 2 years prior to the Great Recession. Mortgage loans, however, have a much longer maturity, about 27 years as shown by the Federal Housing Finance Agency's Monthly Interest Rate Survey.

Entrepreneurs use land as a collateral asset and they borrow at the short rate. Their maximum amount of borrowing is bounded by a fixed portion m of the expected present value of land. The parameters β and ϵ capture entrepreneurs' discount factor and habit formation in consumption. Housing depreciates at a constant rate δ. The depreciation rates of capital are δ_{kc} and δ_{kh} in the consumption and housing sectors, respectively. The last two terms in Eq. (5) capture the capital adjustment costs in consumption and housing sectors; ϕ_{kc} and ϕ_{kh} are the capital adjustment cost parameters.

The production functions in the consumption and housing sectors follow the Cobb–Douglas form:

$$Y_t = \left(a_{c,t}(n'_{c,t})^{\alpha}(n''_{c,t})^{1-\alpha}\right)^{1-\mu_c}(k_{c,t-1})^{\mu_c}, \tag{7}$$

$$IH_t = \left(a_{h,t}(n'_{h,t})^{\alpha}(n''_{h,t})^{1-\alpha}\right)^{1-\mu_h-\mu_b-\mu_l}(k_{h,t-1})^{\mu_h}k_{b,t}^{\mu_b}l_{t-1}^{\mu_l}, \tag{8}$$

where a_c and a_h measure the productivity in the consumption and housing sectors and both follow an AR(1) process. The parameter α measures the labor income share of unconstrained households. Input share parameters are μ_c, μ_h, μ_b, and μ_l.

Technology processes a_c, a_h, and a_k are modeled as

$$\ln a_{c,t} = \rho_c \ln a_{c,t-1} + \epsilon_{c,t}, \tag{9}$$

$$\ln a_{h,t} = \rho_h \ln a_{h,t-1} + \epsilon_{h,t}, \tag{10}$$

$$\ln a_{k,t} = \rho_k \ln a_{k,t-1} + \epsilon_{k,t}, \tag{11}$$

where $\epsilon_{c,t}$, $\epsilon_{h,t}$, and $\epsilon_{k,t}$ are independently and identically distributed innovations with mean zero and variances σ_c^2, σ_h^2, and σ_k^2.

The inter-temporal preference shock z_t follows

$$\ln z_t = \rho_z \ln z_{t-1} + \epsilon_{z,t}, \tag{12}$$

where ρ_z is the autoregressive parameter and $\epsilon_{z,t}$ is an independently and identically distributed innovation with mean zero and variance σ_z^2.

2.2. Households

There is a continuum of measure one of agents in each of the credit-unconstrained and credit-constrained groups. Unconstrained households maximize their lifetime utility

$$V' = E_0 \sum_{t=0}^{\infty}(\beta')^t z_t \left[\frac{1-\epsilon'}{1-\beta'\epsilon'}\ln(c'_t - \epsilon' c'_{t-1}) + j_t \ln(h'_t) - \tau_t\left(n'_{c,t} + n'_{h,t}\right)\right], \tag{13}$$

subject to the following budget constraint:

$$c'_t + q_t(h'_t - (1-\delta)h'_{t-1}) + \frac{R_{S,t-1}b'_{t-1}}{\pi_t} = \frac{w'_{c,t}n'_{c,t}}{X_{wc,t}} + \frac{w'_{h,t}n'_{h,t}}{X_{wh,t}} + DIV'_t + b'_t, \tag{14}$$

where the lump-sum dividends received from labor unions are:

$$DIV'_t = \frac{X_{wc,t}-1}{X_{wc,t}}w'_{c,t}n'_{c,t} + \frac{X_{wh,t}-1}{X_{wh,t}}w'_{h,t}n'_{h,t}. \tag{15}$$

The terms $X_{wc,t}$ and $X_{wh,t}$ denote wage markups. The variables j_t and τ_t capture the housing preference shock and the labor supply shock which follow

$$\ln \tau_t = \rho_\tau \ln \tau_{t-1} + \epsilon_{\tau,t}, \tag{16}$$

$$\ln j_t = (1-\rho_j)\ln j + \rho_j \ln\ln j_{t-1} + \epsilon_{j,t}, \tag{17}$$

where ρ_τ and ρ_j are autoregressive parameters; j is the steady-state value of the housing preference shock; $\epsilon_{\tau,t}$ and $\epsilon_{j,t}$ are independently and identically distributed innovations with mean zero and variances σ_τ^2 and σ_j^2.

The lifetime utility function of the constrained households is

$$V'' = E_0 \sum_{t=0}^{\infty}(\beta'')^t z_t \left[\frac{1-\epsilon''}{1-\beta''\epsilon''}\ln(c''_t - \epsilon'' c''_{t-1}) + j_t \ln(h''_t) - \tau_t\left(n''_{c,t} + n''_{h,t}\right)\right], \tag{18}$$

The constrained households' budget constraint is:

$$c''_t + q_t((h''_t - (1-\delta)h''_{t-1}) + \frac{R_{L,t-1}b''_{t-1}}{\pi_t} = \frac{w''_{c,t}n''_{c,t}}{X_{wc,t}} + \frac{w''_{h,t}n''_{h,t}}{X_{wh,t}} + DIV''_t + b''_t, \tag{19}$$

where the dividends received from labor unions are:

$$DIV''_t = \frac{X_{wc,t}-1}{X_{wc,t}}w''_{c,t}n''_{c,t} + \frac{X_{wh,t}-1}{X_{wh,t}}w''_{h,t}n''_{h,t}. \tag{20}$$

Constrained households use their housing stock as collateral assets and they borrow at the long rate. The borrowing constraint is:

$$b_t'' \leq m'' E_t\left(\frac{q_{t+1}(1-\delta)h_t''\pi_{t+1}}{R_{L,t}}\right), \tag{21}$$

where m'' is the loan-to-value (LTV) ratio associated with credit-constrained households.

2.3. Wage and price stickiness

We introduce Calvo (1983)-type nominal wage stickiness. There are two labor unions in each sector, acting in the interests of two types of households, respectively. Labor unions transform homogeneous labor services that households supply into differentiated labor services. In the meantime, they set nominal wages subject to a Calvo scheme. Each period a fraction of nominal wages cannot be re-optimized and they are indexed to past price inflation instead. The differentiated labor services are then assembled by intermediate labor packers and supplied to wholesale firms. This yields the following four wage Phillips curves:

$$\ln\omega_{c,t}' - \iota_{wc}\ln\pi_{t-1} = \beta'(E_t\ln\omega_{c,t+1}' - \iota_{wc}\ln\pi_t) - \frac{(1-\theta_{wc})(1-\beta'\theta_{wc})}{\theta_{wc}}\ln\left(\frac{X_{wc,t}'}{X_{wc}}\right), \tag{22}$$

$$\ln\omega_{c,t}'' - \iota_{wc}\ln\pi_{t-1} = \beta''(E_t\ln\omega_{c,t+1}'' - \iota_{wc}\ln\pi_t) - \frac{(1-\theta_{wc})(1-\beta''\theta_{wc})}{\theta_{wc}}\ln\left(\frac{X_{wc,t}''}{X_{wc}}\right), \tag{23}$$

$$\ln\omega_{h,t}' - \iota_{wh}\ln\pi_{t-1} = \beta'(E_t\ln\omega_{h,t+1}' - \iota_{wh}\ln\pi_t) - \frac{(1-\theta_{wh})(1-\beta'\theta_{wh})}{\theta_{wh}}\ln\left(\frac{X_{wh,t}'}{X_{wh}}\right), \tag{24}$$

$$\ln\omega_{h,t}'' - \iota_{wh}\ln\pi_{t-1} = \beta''(E_t\ln\omega_{h,t+1}'' - \iota_{wh}\ln\pi_t) - \frac{(1-\theta_{wh})(1-\beta''\theta_{wh})}{\theta_{wh}}\ln\left(\frac{X_{wh,t}''}{X_{wh}}\right), \tag{25}$$

with $\omega_{i,t}$ nominal wage inflation, that is $\omega_{i,t} = \pi_t w_{i,t}/w_{i,t-1}$ for each sector/household pair. The parameters θ_{wc} and θ_{wh} denote the fraction of wage contracts, in consumption and housing sectors, that cannot be re-optimized in each period; ι_{wc} and ι_{wh} capture the degree of wage indexation.

Price stickiness is introduced similarly by assuming monopolistic competition at the retail level, implicit costs of adjusting nominal prices with Calvo-style contracts, and partial indexation to past inflation of those prices that cannot be re-optimized. The resulting inflation equation is:

$$\ln\pi_t - \iota_\pi\ln\pi_{t-1} = \beta'(E_t\ln\pi_{t+1} - \iota_\pi\ln\pi_t) - \frac{(1-\theta_\pi)(1-\beta'\theta_\pi)}{\theta_\pi}\ln\left(\frac{X_t}{X}\right) + \epsilon_{p,t}, \tag{26}$$

where $\epsilon_{p,t}$ is an independently and identically distributed innovation with mean zero and variance σ_p^2. Each period, a fraction $1-\theta_\pi$ of retailers set prices optimally and a fraction θ_π cannot do so. The degree of price indexation is captured by ι_π.

2.4. Taylor rule and the yield curve

We assume that the central bank sets the policy rate $r_{s,t} = \ln(R_{S,t})$ according to a Taylor rule that responds to inflation and GDP growth:

$$r_{s,t} = \gamma_R r_{s,t-1} + (1-\gamma_R)\gamma_\pi\ln\pi_t + (1-\gamma_R)\gamma_Y\ln\left(\frac{GDP_t}{GDP_{t-1}}\right) + (1-\gamma_R)r_s + \eta_t + \epsilon_{rs,t}, \tag{27}$$

where r_s is the steady-state policy rate and $\epsilon_{rs,t}$ is an independently and identically distributed innovation specific to the short-term (1-period) interest rate with mean zero and variance σ_{rs}^2. The term η_t is a persistent AR(1) shock, $\eta_t = \rho_\eta\eta_{t-1} + \epsilon_{\eta,t}$, that captures the central bank's inflation target, where $\epsilon_{\eta,t}$ is an independently and identically distributed innovation with mean zero and variance σ_η^2. GDP is defined as the sum of consumption, residential investment, and non-residential investment weighted by their steady-state nominal shares.

We incorporate the yield curve following De Graeve, Emiris, and Wouters (2009). Under the expectations hypothesis, long-term interest rates reflect economic agents' beliefs about the future path of the short-term rate or policy rate:

$$\hat{r}_{j,t} = \frac{1}{j}E_t(\hat{r}_{s,t} + \hat{r}_{s,t+1} + \cdots + \hat{r}_{s,t+j-1}) \text{ for } j = 2, 3, \ldots, l, \tag{28}$$

where $\hat{r}_{j,t}$ denotes the log deviation of the j-period interest rate from its steady state.[3] Its observed counterpart is

$$r_{j,t} = \hat{r}_{j,t} + r_s + c_j, \tag{29}$$

[3] Kulish et al. (2017) augment the structural model of Smets and Wouters (2007) with a similar yield curve specification. It is worth noting that our specification of the yield curve is slightly different from that of Kulish, Morley, and Robinson (2017). In their paper, the authors did not take into account the persistent inflation target, η_t, in the policy rate but additively built it into all other interest rates as part of the measurement errors for the purpose of estimation. Interest rate fluctuations at different maturities are exogenously determined in their model, while we allow them to be driven by economic agents' expectations by building the inflation target into the policy rate. Another stand of literature incorporates the yield curve into DSGE models from the asset pricing perspective; see Emiris (2006), Van Binsbergen et al. (2012), and Vázquez and Aguilar (2021), among others.

where $r_s + c_j$ is the steady state of the j-period interest rate and c_j is the difference between the j-period interest rate and the policy rate. Adding shocks (or measurement errors) for the purpose of estimation gives

$$\hat{r}_{j,t} = \frac{1}{j} E_t(\hat{r}_{s,t} + \hat{r}_{s,t+1} + \cdots + \hat{r}_{s,t+j-1}) + \epsilon_{rj,t} \text{ for } j = 2, 3, \ldots, l, \tag{30}$$

where $\epsilon_{rj,t}$ is an independently and identically distributed innovation specific to the j-period interest rate with mean zero and variance σ^2_{rj}. While these measurement errors are added to improve the fit of the expectations hypothesis, they also capture the discrepancy between the expectations-hypothesis implied yield and the observed yield, which can be interpreted as a measure of fluctuations in the term premium; see De Graeve et al. (2009).

2.5. Market clearing conditions

The market clearing conditions are:

$$C_t + IK_{c,t}/a_{k,t} + IK_{h,t} + k_{b,t} = Y_t - \Phi_t, \tag{31}$$

$$(h'_t + h''_t) - (1-\delta)(h'_{t-1} + h''_{t-1}) = IH_t, \tag{32}$$

$$b_t + b'_t + b''_t = 0, \tag{33}$$

where Φ_t is the sum of capital adjustment costs in both sectors.

3. Parameter estimates

3.1. Data description

Our sample for the estimation of the DSGE model ranges from 1972:Q1 to 2019:Q4, which lines up with what has been used in the introduction section. We use data for 13 variables: real consumption, real business investment, real residential investment, real house prices, inflation rate, hours and wage inflation in the consumption sector, hours and wage inflation in the housing sector, and 3-month, 2-year, 5-year, and 10-year interest rates. The 3-month and 10-year interest rates correspond to the short- and long-term interest rates in our model, respectively. We also include the 2-year and 5-year interest rates to better approximate the shape of the entire yield curve. Linear trends are removed prior to estimation.

Real consumption, real business investment, and real residential investment are obtained from the Bureau of Economic Analysis (BEA), expressed in per-capita terms; inflation rate is the percent change of the implicit price deflator in nonfarm business sector; real house prices are price indexes of new single family houses sold including lot value obtained from the U.S. Census Bureau, deflated with the implicit price deflator; hours and wage inflations are obtained from the U.S. Bureau of Labor Statistics (BLS); the 3-month Treasury Bill Rate at secondary market is obtained from the Board of Governors of the Federal Reserve System; 2-year, 5-year and 10-year interest rates are zero-coupon yields calculated by Gürkaynak, Sack, and Wright (2007).

The 3-month Treasury Bill Rate at secondary market stays near the zero lower bound in the aftermath of the 2018 financial crisis, which makes the movement at the short end of the yield curve irrelevant. In order to capture the effects of unconventional monetary policy, we add the difference between the shadow rate of Wu and Xia (2016) and the federal funds rate to the 3-month Treasury Bill Rate to measure the short-term interest rate after 2008. The Wu and Xia (2016) shadow rate has been used in the estimation of New Keynesian models by many recent papers, e.g., Wu and Zhang (2019), Mouabbi and Sahuc (2019), and Aguirre and Vázquez (2020).

3.2. Calibration

We calibrate the discount factors (β, β' and β''), the production function parameters (μ_c, μ_h, μ_b and μ_l), the depreciation rates (δ, δ_{kc} and δ_{kh}), the LTV ratios (m and m''), the steady-state value of the housing preference shock (j), and the steady-state gross price and wage markups (X, X_{wc} and X_{wh}) as in Iacoviello and Neri (2010); see Table 2. These values match the shares of consumption, business investment, and housing investment in GDP of around 67%, 27%, and 6%, respectively. The steady-state term spreads (C_8, C_{20} and C_{40}) are calibrated to match the average values of the 2-year, 5-year, and 10-year interest rates less the 3-month short-term interest rate.

The prior distributions of the structural parameters and shock processes are presented in the left panel of Table 3. These choices largely follow Iacoviello and Neri (2010), Sun and Tsang (2017), and Smets and Wouters (2007). The right panel reports the posterior mean, median, and 90% probability intervals. The estimated labor income share of credit-unconstrained households α is 0.78, which implies a share of credit-constrained households of about 22%. The Taylor rule parameter estimates are consistent with previous evidence. Nominal price and wage rigidities (θ_π, θ_{wc} and θ_{wh}) are estimated to be quite strong. All AR(1) shocks are estimated to be highly persistent as in Iacoviello and Neri (2010).

4. Properties of the estimated model

In this section, we first examine the model's performance in capturing the yield spread fluctuations and present the impulse response functions of two interest rate shocks specific to the short and long ends of the yield curve. Then, through a series of counterfactual analyses, we try to identify key mechanisms that determine the predictive power of yield spread for future economic activities.

Table 2
Calibrated parameters.

Parameter	Value	Parameter	Value	Parameter	Value
β	0.9925	μ_l	0.20	j	0.18
β'	0.9943	δ	0.012	X	1.15
β''	0.95	δ_{kc}	0.032	X_{wc}, X_{wh}	1.15
μ_c	0.40	δ_{kh}	0.035	C_8	0.0023
μ_h	0.10	m	0.60	C_{20}	0.0035
μ_b	0.10	m''	0.85	C_{40}	0.0048

Table 3
Prior and posterior distribution of model parameters.

Parameter	Prior distribution			Posterior distribution			
	Distribution	Mean	SD	Mean	5%	Median	95%
α	Beta	0.65	0.05	0.7816	0.7564	0.7850	0.7916
ϵ	Beta	0.5	0.075	0.6194	0.6069	0.6163	0.6372
ϵ'	Beta	0.5	0.075	0.5693	0.5607	0.5651	0.5963
ϵ''	Beta	0.5	0.075	0.6884	0.6742	0.6847	0.7150
ι_π	Beta	0.5	0.2	0.6712	0.6102	0.6846	0.6914
ι_{wc}	Beta	0.5	0.2	0.1299	0.1146	0.1223	0.1704
ι_{wh}	Beta	0.5	0.2	0.4384	0.4144	0.4302	0.4845
γ_R	Beta	0.75	0.05	0.6808	0.6563	0.6862	0.7035
γ_π	Normal	1.5	0.05	1.3501	1.3262	1.3296	1.4249
γ_Y	Normal	0	0.05	0.0938	0.0810	0.0877	0.1164
θ_π	Beta	0.667	0.1	0.9450	0.9402	0.9449	0.9506
θ_{wc}	Beta	0.667	0.1	0.9164	0.8730	0.9261	0.9329
θ_{wh}	Beta	0.667	0.1	0.9425	0.8611	0.9604	0.9668
ϕ_{kc}	Gamma	10	2.5	4.1752	4.1690	4.1731	4.1895
ϕ_{kh}	Gamma	10	2.5	9.2782	9.2504	9.2610	9.3495
ρ_{AC}	Beta	0.8	0.1	0.9576	0.9296	0.9672	0.9714
ρ_{AH}	Beta	0.8	0.1	0.9323	0.9207	0.9252	0.9578
ρ_{AK}	Beta	0.8	0.1	0.9980	0.9835	0.9995	0.9998
ρ_j	Beta	0.8	0.1	0.9344	0.9272	0.9331	0.9463
ρ_τ	Beta	0.8	0.1	0.8402	0.8312	0.8359	0.8584
ρ_z	Beta	0.8	0.1	0.9877	0.9855	0.9877	0.9920
ρ_η	Beta	0.8	0.1	0.9551	0.9201	0.9635	0.9668
σ_{AC}	Inv. Gamma	0.001	0.02	0.0145	0.0127	0.0141	0.0180
σ_{AH}	Inv. Gamma	0.001	0.02	0.0289	0.0243	0.0295	0.0320
σ_{AK}	Inv. Gamma	0.001	0.02	0.0247	0.0209	0.0241	0.0320
σ_j	Inv. Gamma	0.001	0.02	0.1054	0.0823	0.1084	0.1108
σ_τ	Inv. Gamma	0.001	0.02	0.0727	0.0482	0.0765	0.0848
σ_z	Inv. Gamma	0.001	0.02	0.0291	0.0237	0.0294	0.0343
σ_η	Inv. Gamma	0.001	0.02	0.0004	0.0003	0.0003	0.0007
σ_{rs}	Inv. Gamma	0.001	0.02	0.0030	0.0027	0.0030	0.0034
σ_{rl}	Inv. Gamma	0.001	0.02	0.0014	0.0011	0.0013	0.0026
σ_p	Inv. Gamma	0.001	0.02	0.0041	0.0038	0.0041	0.0045
σ_{NH}	Inv. Gamma	0.001	0.02	0.1366	0.1329	0.1352	0.1431
σ_{WH}	Inv. Gamma	0.001	0.02	0.0045	0.0041	0.0045	0.0049
σ_{r8}	Inv. Gamma	0.001	0.02	0.0012	0.0011	0.0012	0.0016
σ_{r20}	Inv. Gamma	0.001	0.02	0.0007	0.0005	0.0006	0.0013

In order to ensure that the number of shocks is more than or equal to the number of data series, we impose independently and identically distributed measurement errors on hours and wage inflation in the housing sector, with their standard errors being denoted σ_{NH} and σ_{WH}.

4.1. Yield spread fluctuations

In order to verify that our model is able to capture most of the fluctuations in the yield spread, we take the estimated model as given and conduct a counterfactual analysis by simulating model variables from the estimated parameters and shocks with the shock (or measurement error) specific to the long-term interest rate, ϵ_{lt}, shut off. Fig. 1 presents the counterfactual yield spread with a red dashed line and its observed counterpart with a blue solid line. As the figure shows, by building the inflation target into the policy rate and allowing it to affect the long-term interest rate through economic agents' rational expectations, our model is able to explain most of the observed variation in yield spread without relying on measurement errors.

To better understand the driving forces of the yield curve at both ends, we present in Table 4 the results of variance decomposition for the short- and long-term interest rates at business cycle frequencies. Structural shocks that drive the movements at the short and long ends of the yield curve are different. The variance in the short-term interest rate is dominantly explained by the cost-push shock and the short-term interest rate shock, which account for about one half and one third of the variation at the short end of

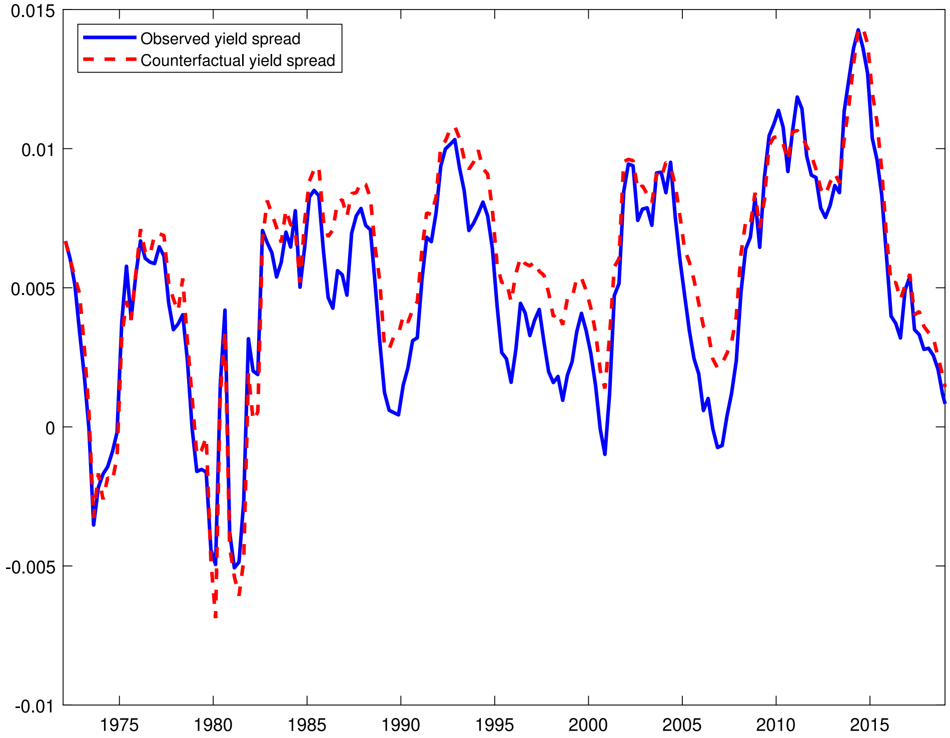

Fig. 1. Observed and counterfactual spread between 10-year and 3-month interest rates.

Table 4
Variance decomposition of interest rates.

	Short-term interest rate	Long-term interest rate
Consumption sector productivity shock ϵ_c	3.47%	11.13%
Housing sector productivity shock ϵ_h	0.07%	0.03%
Investment-specific shock ϵ_k	2.52%	14.20%
Housing preference shock ϵ_j	0.21%	0.04%
Labor supply shock ϵ_τ	1.17%	5.05%
Inter-temporal preference shock ϵ_z	4.36%	35.47%
Inflation target shock ϵ_η	1.11%	6.77%
Short-term interest rate shock ϵ_{rs}	34.31%	2.15%
Long-term interest rate shock ϵ_{rl}	0.00%	21.23%
Cost-push shock ϵ_p	52.78%	3.93%

the yield curve, respectively. The fluctuations of the long-term interest rate are mainly explained by the inter-temporal preference shock (about 35%), the long-term interest rate shock (about 20%), and the investment-specific shock (about 15%). Given that the expectation hypothesis models the long-term interest rate as an average of expected future short-term interest rates, it does not come as a surprise that the inter-temporal preference shock contributes most to the fluctuations of the long-term interest rate.

4.2. Impulse responses

We are interested in the movements at the two ends of the yield curve and how they affect future economic activities. We plot the impulse response functions of eight key model variables, i.e., real consumption, real business investment, real housing investment, real house prices, price inflation, short- and long-term interest rates, and real GDP, to the shocks specific to the short- and long-term interest rates. The steady-state spread between long-term and short-term interest rates is 0.48 percentage points. We find that either a positive shock of 2 standard deviations at the short end ($\epsilon_{rs,t}$) of the yield curve or a negative shock of 3.4 standard deviations at the long end ($\epsilon_{rl,t}$) is just enough to lead to an inversion of the yield curve from the steady state. We plot the impacts of such shocks at the two ends of the yield curve in Fig. 2.

Following a positive shock to the short-term interest rate, all components of aggregate demand fall, with business investment showing the largest drop, followed by housing investment and consumption. This finding is different from Iacoviello and Neri (2010), where the largest drop following a monetary policy shock is on housing investment. This difference results from the fact that the demand side of the housing market is affected by the long-term interest rate rather than the policy rate as in their model. Such a

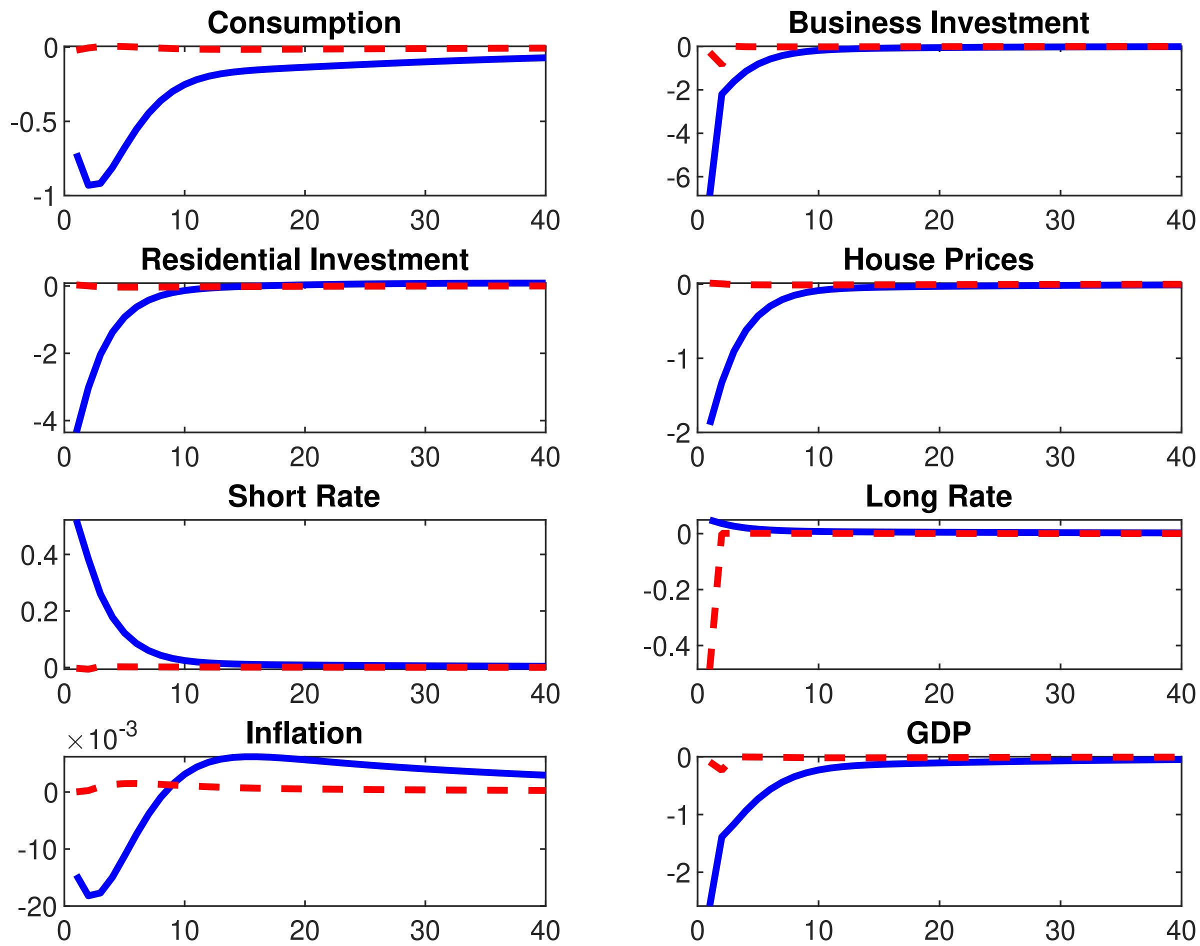

Fig. 2. Impulse responses to interest rate shocks.
Note: The *y*-axis measures the percent deviation from the steady state. The blue solid line shows the response to a positive short-term interest rate shock of 2 standard deviations and the dashed red line shows the response to a negative long-term interest rate shock of 3.4 standard deviations. (For interpretation of the references to color in this figure legend, the reader is referred to the web version of this article.)

shock that is enough to invert the yield curve causes a 2 percent decrease in real GDP. The negative impact on real GDP gradually diminishes and lasts for about 10 quarters.

Following a negative shock to the long-term interest rate, both consumption and business investment decrease. Residential investment and house prices increase during the first few quarters and then decrease. Real GDP decreases. However, the negative impact on real GDP, about 0.3 percent at the maximum, is much smaller in size than that caused by the shock at the short end of the yield curve. This figure clearly shows that an increase in the short-term interest rate and a decrease in the long-term interest rate have different impacts on the economy. This result supports the argument of Wright (2006) that there is no fundamental reason why movements at the two ends of the yield curve must have the same predictive content for the likelihood of a recession.

4.3. Counterfactual analyses

4.3.1. Yield spread as predictor of future recessions

We simulate the short- and long-term interest rates and real GDP from the estimated model parameters and structural shocks, and use the simulated yield spread to predict future recessions. Given that we use simulated data, a recession here is defined as a period of declining (simulated) real GDP over two consecutive quarters. We estimate the same set of probit regression models over a four-quarter horizon as in the introduction.

Table 5 displays the results from the probit model estimates over three subsamples, 1972:Q1–1990:Q3, 1990:Q4–2007:Q1, and 2007:Q2–2019:Q4. The splitting of the sample follows Table 1 in the introduction. Over the first subsample period, the yield spread itself is a statistically significant predictor of future recessions. As the yield curve becomes flatter (or the yield spread decreases), the probability of having a recession over the next four quarters increases. The yield spread remains significant in both regressions with either the short rate or the long rate being controlled. However, its coefficient is larger in magnitude when the long rate is

Table 5
Yield spread as predictor of future recessions — benchmark model.

	1975:Q1-1990:Q3			1990:Q4–2007:Q1			2007:Q2–2019:Q4		
Spread	−186.58***	−179.03**	−186.69***	−149.77	−183.90*	−144.27	−161.80**	−222.56	−159.47**
	(−2.71)	(−2.13)	(−2.72)	(−1.58)	(−1.67)	(−1.50)	(−2.05)	(−1.53)	(−2.09)
Short rate		7.67			−39.62			−63.09	
		(0.21)			(−0.41)			(−0.58)	
Long rate			7.67			−39.62			−63.09
			(0.21)			(−0.41)			(−0.58)
N	75	75	75	66	66	66	49	49	49
Mc Fadden R^2	0.178	0.178	0.178	0.099	0.105	0.105	0.154	0.161	0.161

Entries in parentheses are t statistics constructed using Newey–West standard errors. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table 6
Yield spread as predictor of future recessions — the impact of price and wage rigidities.

	Benchmark		$\theta_\pi = 0.5$		$\theta_{wc} = 0.5$		$\theta_{wh} = 0.5$	
Spread	−149.77	−183.90*	4.30	−125.60	−102.41	−51.42	−106.18	−55.83
	(−1.58)	(−1.67)	(0.09)	(−1.04)	(−1.54)	(−0.46)	(−1.28)	(−0.58)
Short rate		−39.62		−95.61		56.27		57.78
		(−0.41)		(−1.09)		(0.56)		(0.69)
N	66	66	66	66	66	66	66	66
Mc Fadden R^2	0.099	0.105	0.001	0.034	0.062	0.067	0.053	0.064

Results are obtained from probit regressions using simulated data over the second subsample period 1990:Q4–2007:Q1. Entries in parentheses are t statistics constructed using Newey–West standard errors. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

controlled. This result implies that a decrease in the yield spread arising from a rise in the short-term interest rate predicts a higher probability of future recessions over the 1972:Q1–1990:Q3 period. Over the second subsample period 1990:Q4–2007:Q1, the yield spread is generally not a significant predictor of future recessions but it significantly predicts future recessions when the short rate is controlled, and a decrease in the yield spread arising from a fall in the long-term interest rate predicts a higher probability of future recessions. Over the third subsample period 2007:Q2–2019:Q4, movements at the short end of the yield curve regain predictive power, as the yield spread stays significant in the regression which controls for the long-term interest rate but becomes insignificant in the regression which controls for the short-term interest rate instead. These results are line with the results presented in Table 1 using recessions defined by NBER.

It is not surprising that movements at the long end of the yield curve better predict future recessions over the second subsample period which covers the housing boom started in mid-to-late 1990s. What affects the predictive power of the yield spread at the long end, however, is far from clear. One possibility is that movements at the long end of the yield curve better predict future recessions as the sectors heavily affected by long-term interest rates, such as the housing sector, experience more volatile fluctuations, captured by the structural shocks in the DSGE model. However, it is also possible that some key structural parameters of the model that capture the underlying mechanisms behind the macroeconomy determine how long-term interest rates affect future economic activities.

Next, through a series of counterfactual analyses, we would like to identify a few model parameters that might affect the magnitude of the impact of the long-term interest rate on future economic activities. We focus our attention on the second subsample period over which the housing sector plays an important role. We find that most of the structural model parameters either do not affect the predictive power of the yield spread at all or affect it to a minor extent without any clear pattern. Estimated parameters that have a significant impact on the predictive power of the yield spread at the long end include price and wage rigidity parameters (θ_π, θ_{wc}, and θ_{wh}) and the capital adjustment cost in the consumption sector (ϕ_{kc}).

4.3.2. Nominal price and wage rigidities

In our benchmark model, the price rigidity parameter θ_π is estimated to be 0.9450, and the wage rigidity parameters are estimated to be $\theta_{wc} = 0.9164$ and $\theta_{wh} = 0.9425$ in the consumption and housing sectors, respectively. In other words, estimation suggests that 90% of the price and wage contracts cannot be re-optimized in each period. We assume more flexible prices or wages in the counterfactual analysis by setting the three rigidity parameters, one at a time, to a lower value of 0.5. Other model parameters take on their estimated values as reported in Table 3. Given the estimated shocks, we simulate short- and long-term interest rates and real GDP, and then use the simulated yield spread to predict recessions over the next four quarters over the second subsample period 1990:Q4–2007:Q1. This allows us to identify how each of these parameters affects the predictive power of movements at the long end of the yield curve.

As Table 6 shows, both price and wage rigidities affect the predictive power of the yield spread. With lower nominal rigidities (i.e., $\theta_\pi = 0.5$, $\theta_{wc} = 0.5$, or $\theta_{wh} = 0.5$), the statistical significance of the yield spread disappears, even when the short rate is controlled in the probit regression. The goodness of fit of the model, measured by the Mc Fadden R^2, deteriorates. The intuition is straightforward. The long-term interest rate affects the macroeconomy mainly through its impact on housing demand, but real variables, especially real housing investment, become less sensitive to structural shocks when nominal prices or wages are more flexible.

Table 7
Yield spread as predictor of future recessions — the impact of capital adjustment costs.

	Benchmark		$\phi_{kc}=0$		$\phi_{kc}=20$	
Spread	−149.77	−183.90*	−148.81	−187.10*	−84.38	−93.13
	(−1.58)	(−1.67)	(−1.62)	(−1.69)	(−1.05)	(−0.92)
Short rate		−39.62		−44.27		−10.41
		(−0.41)		(−0.46)		(−0.13)
N	66	66	66	66	66	66
Mc Fadden R^2	0.099	0.105	0.101	0.108	0.030	0.031

Results are obtained from probit regressions using simulated data over the second subsample period 1990:Q4–2007:Q1. Entries in parentheses are t statistics constructed using Newey–West standard errors. $^*p<0.10$, $^{**}p<0.05$, $^{***}p<0.01$.

4.3.3. Capital adjustment costs

The capital adjustment cost affects the responsiveness of investment to structural shocks. As in Iacoviello and Neri (2010), a version of the model without capital adjustment costs exacerbates the volatility of both residential and nonresidential investment relative to consumption. Our estimated value of ϕ_{kc} is 4.1752; see Table 3. In the counterfactual analysis, we assume that the capital adjustment cost parameter in the consumption sector either takes a smaller value of 0 or a larger value of 20. As Table 7 shows, when we assume $\phi_{kc}=0$, real housing investment exhibits more variation, and therefore movements at the long end of the yield curve have a larger impact on the likelihood of future recessions. The goodness of fit of the regression slightly improves. When we assume higher capital adjustment costs, e.g., $\phi_{kc}=20$, the yield spread loses its statistical significance at the long end. The R^2 reduces sharply to about 3%.

5. A regime-switching version of the model

In the previous section, through a series of counterfactual analyses, we identify a set of key model parameters that affect the predictive power of the yield spread, i.e., price and wage rigidities parameters (θ_π, θ_{wc}, and θ_{wh}) and one of the capital adjustment cost parameters (ϕ_{kc}). We find that, taking the estimated structural shocks as given, the predictive power of the yield curve at the long end improves when nominal rigidities increase or the capital adjustment cost decreases. Given that this finding matches the pattern identified from actual data presented in Table 1 over the three subsample periods, it is straightforward to suspect that the fact that movements at the long end of the yield curve better predict future recessions over the second subsample period might be driven by changes in these key parameters. While we assume that our structural model parameters take on constant values over the entire sample in previous sections, we now introduce regime switching to both wage rigidity parameters (θ_{wc} and θ_{wh}) and both capital adjustment cost parameters (ϕ_{kc} and ϕ_{kh}).[4] The model features two regimes, the one with lower wage rigidities is named regime 1 and the other regime 2. We do not impose any restrictions on the capital adjustment cost parameters. These four parameters are allowed to switch as a Markov chain with the transition matrix

$$Q=\begin{bmatrix} 1-p_{1,2} & p_{1,2} \\ p_{2,1} & 1-p_{2,1} \end{bmatrix}, \tag{34}$$

where $p_{1,2}$ is the probability of switching from regime 1 in period t to regime 2 in period $t+1$, $1-p_{1,2}$ is the probability of remaining in regime 1, $p_{2,1}$ is the probability of switching from regime 2 in period t to regime 1 in period $t+1$; and $1-p_{2,1}$ is the probability of remaining in regime 2. All other parameters are assumed to be constant over the entire sample period. We solve the regime switching DSGE model with the perturbation method proposed by Maih (2015),[5] and the switching parameter estimates and the transition probabilities are presented in Table 8. The prior distributions of the model parameters stay the same as in Section 2 and those of switching probabilities are taken from Maih (2015).

The wage rigidity parameters are estimated to be $(\theta_{wc},\theta_{wh})=(0.7886,0.7945)$ in regime 1 and $(\theta_{wc},\theta_{wh})=(0.9177,0.9396)$ in regime 2. In the meantime, the capital adjustment cost parameter estimates are $(\phi_{kc},\phi_{kh})=(6.2647,15.2889)$ in regime 1 and $(\phi_{kc},\phi_{kh})=(2.7362,6.5569)$ in regime 2. Fig. 3 plots the smoothed probabilities of regime 2, where wage rigidities are estimated to be higher and capital adjustment costs are estimated to be lower. Except for the first three years, the economy stays in regime 1 that features lower wage rigidities and higher capital adjustment costs over the early sample period. The transition to regime 2 with higher wage rigidities and lower capital adjustment costs occurs around the year 1990. Since then, the economy has remained in regime 2 until it switches back to regime 1 around 2007 and 2008 when the housing market collapses. By introducing regime switching into the model and allowing key model parameters to switch between two regimes, we confirm our conjecture that, compared to the first and third subsample periods, the second subsample period, over which movements at the long end of the yield curve outweigh those at the short end in predicting future recessions, features higher wage rigidities and lower capital adjustment costs.

[4] We also allow the price rigidity parameter θ_π to switch between two regimes but do not find such a change.

[5] The perturbation method of Maih (2015) has been adopted by Bjørnland et al. (2018), Jin and Xiong (2021), Maih et al. (2021), Chang et al. (2021), and Benchimol and Ivashchenko (2021) among others.

Table 8
Estimates of regime-switching parameters and transition probabilities.

Parameter	Prior distribution			Posterior distribution			
	Distribution	Mean	SD	Mean	5%	Median	95%
$\theta_{wc}(1)$	Beta	0.667	0.1	0.7886	0.7731	0.7908	0.7935
$\theta_{wc}(2)$	Beta	0.667	0.1	0.9177	0.8989	0.9218	0.9274
$\theta_{wh}(1)$	Beta	0.667	0.1	0.7945	0.7853	0.7942	0.7996
$\theta_{wh}(2)$	Beta	0.667	0.1	0.9396	0.9240	0.9379	0.9554
$\phi_{kc}(1)$	Gamma	10	2.5	6.2647	6.2609	6.2640	6.2734
$\phi_{kc}(2)$	Gamma	10	2.5	2.7362	2.7161	2.7393	2.7465
$\phi_{kh}(1)$	Gamma	10	2.5	15.2889	15.2786	15.2915	15.2967
$\phi_{kh}(2)$	Gamma	10	2.5	6.5569	6.5457	6.5593	6.5640
$p_{1,2}$	Beta	0.3333	0.0556	0.0203	0.0100	0.0224	0.0311
$p_{2,1}$	Beta	0.3333	0.0556	0.0395	0.0241	0.0433	0.0485

Estimates of non-switching parameters are not reported.

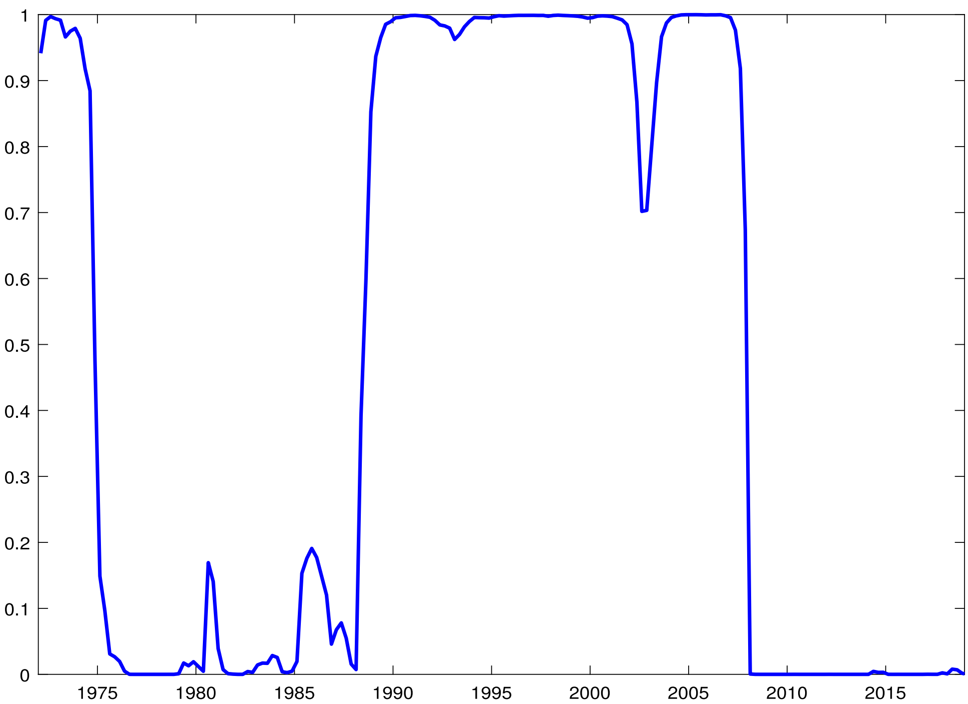

Fig. 3. Smoothed probabilities of regime 2.

Having confirmed the changes in wage rigidities and capital adjustment costs with a regime-switching DSGE model, we are now able to check whether the higher predictive power of the yield curve at the long end over the later sample period is driven by these parameter changes. We take the estimated structural shocks as given and assume that wage rigidities and capital adjustment costs have stayed at their regime-1 values between 1990 and 2008. We then simulate the short- and long-term interest rates and real GDP, and use the simulated yield spread to predict future recessions, defined as periods of declining (simulated) real GDP over two consecutive quarters. We present the counterfactual regression results in the right panel of Table 9, with the benchmark results in the left panel for the sake of comparison. If wage rigidities and capital adjustment costs had stayed at their regime-1 values, the yield spread would lose its predictive power, even when we control for the short-term interest rate, and the goodness of fit of the regressions would deteriorate.

6. Further evidence from actual data

Our counterfactual analyses show that both wage rigidities and capital adjustment costs are among the underlying mechanisms through which the long-term interest rate affects future economic activities. The size of the negative impact of the long rate on future economic activities increases with the degree of wage rigidity and decreases with the capital adjustment cost. While wage

Table 9
Yield spread as predictor of future recessions — the impact of nominal wage rigidities and capital adjustment costs.

	Benchmark		Counterfactual	
Spread	−149.77	−183.90*	−129.90	−124.18
	(−1.58)	(−1.67)	(−1.62)	(−1.12)
Short rate		−39.62		6.37
		(−0.41)		(0.06)
N	66	66	66	66
Mc Fadden R^2	0.099	0.105	0.079	0.079

Results are obtained from probit regressions using simulated data over the second subsample period 1990:Q4–2007:Q1. Entries in parentheses are t statistics constructed using Newey–West standard errors. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

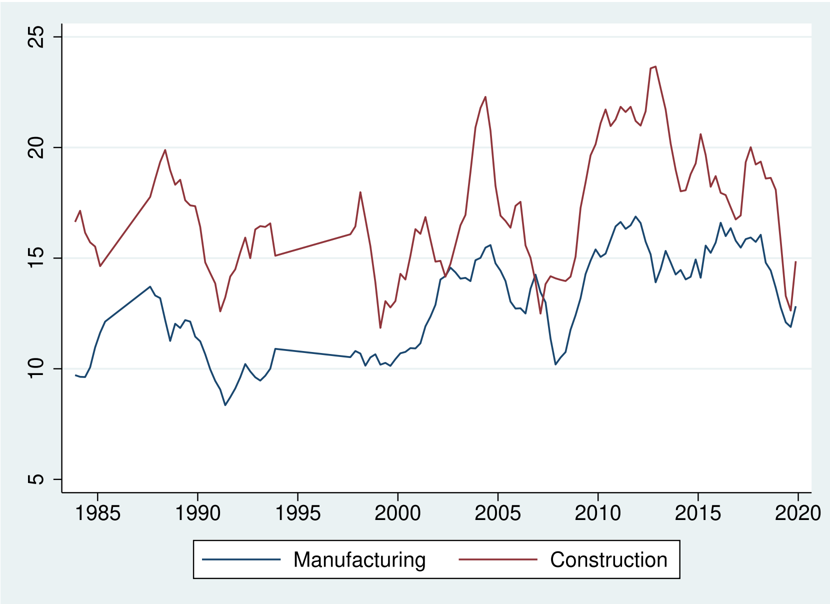

Fig. 4. Nominal wage rigidities by industry.

rigidities are structural parameters in our model, they have observable counterparts though the match is not perfect.[6] In this section, we first examine the behavior of observable wage rigidities over time and then include their interactions with yield spread in the probit regressions as extra predictors for future NBER recessions.

Fig. 4 presents the share of individuals with zero nominal wage change in a particular year relative to the previous year and illustrates how nominal wage rigidities have behaved over time. These data are at monthly frequency and are taken from the Federal Reserve Bank of San Francisco's Wage Rigidity Meter, which is constructed from Current Population Survey (CPS) data on individuals that have not changed jobs over the course of a year; see Daly, Hobijn, and Lucking (2012) and Daly, Hobijn, and Wiles (2011).

With incomplete data in the first subsample period, the share of individuals with zero nominal wage change in the construction industry remains stable on average between the first two subsamples. In the manufacturing sector, however, a larger share of individuals experience zero nominal wage change over the second subsample period. Wage rigidities in both sectors have been increasing over most of the Great Moderation period before they start to decline in 2012, which is sensible given the downward trend in inflation since the 1980s. Kahn (1997) finds that nominal wage adjustments are less likely to occur when a given real wage change requires a small nominal change (i.e., low inflation) than when it requires a larger nominal change (i.e., high inflation). According to Coibion and Gorodnichenko (2015), falling levels of inflation since the 1980s have lowered nominal wage changes. The increasing pattern of nominal wage rigidities over our second subsample period supports the argument that wage rigidity could be one of the mechanisms through which the long rate predicts future economic activities.

Next, we use the yield spread as predictor of actual data recessions, defined by the NBER's Business Cycle Dating Committee, adding its interactions with observable wage rigidities as extra predictors.[7] We present the parameter estimates from the probit

[6] A measurable counterpart of capital adjustment costs is not available.

[7] The monthly data on wage rigidities are converted into quarterly data by taking the un-weighted quarterly average to match the frequency of all other variables.

Table 10
Yield spread as predictor of future NBER recessions augmented with wage rigidities.

Spread	−238.21***	−264.13***	323.52	1310.34*
	(−2.67)	(−3.41)	(1.32)	(1.90)
Short rate		−27.56	−77.42	−67.72
		(−0.83)	(−1.26)	(−1.16)
Spread × Rigidity$_M$			−52.35**	
			(−2.31)	
Spread × Rigidity$_C$				−104.99**
				(−2.28)
N	145	145	145	145
Mc Fadden R^2	0.231	0.240	0.290	0.355

Results are obtained from probit regressions using actual data. Spread × Rigidity$_M$ and Spread × Rigidity$_C$ stand for the interactions of the yield spread with the measurable wage rigidities in the manufacturing and construction sectors. Entries in parentheses are t statistics constructed using Newey–West standard errors. *$p < 0.10$, **$p < 0.05$, ***$p < 0.01$.

regressions in Table 10. The table shows negative and significant coefficients of the interaction between yield spread and wage rigidities. This result clearly suggests that yield spread better predicts future NBER recessions in the case of high wage rigidities, which supports the results of our counterfactual analyses.

7. Conclusion

In this paper, we investigate the changing predictive power of the yield spread with the help of a medium-sized DSGE model. More generally, we believe that structural macroeconomic models are a useful tool to make sense of reduced-form regression results. By changing the parameters of the model and simulating data, we are able to find out which parts of the model are closely related to the empirical regularities and which are not, and it is particularly helpful when we try to explain why some reduced-form results are unstable.

Declaration of competing interest

The authors declare that they have no known competing financial interests or personal relationships that could have appeared to influence the work reported in this paper.

Data availability

Data will be made available on request.

References

Aguirre, I., Vázquez, J., 2020. Learning, parameter variability, and swings in US macroeconomic dynamics. J. Macroecon. 66, 103240.
Bai, J., Perron, P., 1998. Estimating and testing linear models with multiple structural changes. Econometrica 47–78.
Bai, J., Perron, P., 2003. Computation and analysis of multiple structural change models. J. Appl. Econometrics 18 (1), 1–22.
Benchimol, J., Ivashchenko, S., 2021. Switching volatility in a nonlinear open economy. J. Int. Money Finance 110, 102287.
Bjørnland, H.C., Larsen, V.H., Maih, J., 2018. Oil and macroeconomic (in) stability. Am. Econ. J.: Macroecon. 10 (4), 128–151.
Calvo, G.A., 1983. Staggered prices in a utility-maximizing framework. J. Monetary Econ. 12 (3), 383–398.
Chang, Y., Maih, J., Tan, F., 2021. Origins of monetary policy shifts: A new approach to regime switching in DSGE models. J. Econom. Dynam. Control 133, 104235.
Chinn, M., Kucko, K., 2015. The predictive power of the yield curve across countries and time. Int. Finance 18 (2), 129–156.
Coibion, O., Gorodnichenko, Y., 2015. Is the Phillips curve alive and well after all? Inflation expectations and the missing disinflation. Am. Econ. J.: Macroecon. 7 (1), 197–232.
Daly, M., Hobijn, B., Lucking, B., 2012. Why has wage growth stayed strong? FRBSF Econ. Lett. 10 (2).
Daly, M., Hobijn, B., Wiles, T., 2011. Dissecting Aggregate Real Wage Fluctuations: Individual Wage Growth and the Composition Effect. FRBSF Working Paper, Federal Reserve Bank of San Francisco.
De Graeve, F., Emiris, M., Wouters, R., 2009. A structural decomposition of the US yield curve. J. Monetary Econ. 56 (4), 545–559.
Dueker, M.J., 1997. Strengthening the case for the yield curve as a predictor of US recessions. Federal Reserve Bank of St. Louis Review 79 (2), 41.
Emiris, M., 2006. The Term Structure of Interest Rates in a DSGE Model. National Bank of Belgium Working Paper 88.
Estrella, A., Hardouvelis, G.A., 1991. The term structure as a predictor of real economic activity. J. Finance 46 (2), 555–576.
Estrella, A., Mishkin, F.S., 1996. The yield curve as a predictor of US recessions. Curr. Issues Econ. Finance 2 (7).
Estrella, A., Mishkin, F.S., 1998. Predicting US recessions: Financial variables as leading indicators. Rev. Econ. Stat. 80 (1), 45–61.
Gürkaynak, R.S., Sack, B., Wright, J.H., 2007. The US treasury yield curve: 1961 to the present. J. Monetary Econ. 54 (8), 2291–2304.
Iacoviello, M., Neri, S., 2010. Housing market spillovers: Evidence from an estimated DSGE model. Am. Econ. J.: Macroecon. 2 (2), 125–164.
Jin, H., Xiong, C., 2021. Fiscal stress and monetary policy stance in oil-exporting countries. J. Int. Money Finance 111, 102302.
Kahn, S., 1997. Evidence of nominal wage stickiness from microdata. Am. Econ. Rev. 87 (5), 993–1008.
Kulish, M., Morley, J., Robinson, T., 2017. Estimating DSGE models with zero interest rate policy. J. Monetary Econ. 88, 35–49.
Maih, J., 2015. Efficient Perturbation Methods for Solving Regime-Switching DSGE Models. Norges Bank Working Paper 1| 2015.
Maih, J., Mazelis, F., Motto, R., Ristiniemi, A., 2021. Asymmetric monetary policy rules for the Euro area and the US. J. Macroecon. 70, 103376.
Mouabbi, S., Sahuc, J.-G., 2019. Evaluating the macroeconomic effects of the ECB's unconventional monetary policies. J. Money Credit Bank. 51 (4), 831–858.

Smets, F., Wouters, R., 2007. Shocks and frictions in US business cycles: A Bayesian DSGE approach. Amer. Econ. Rev. 97 (3), 586–606.
Stock, J.H., Watson, M.W., 1989. New indexes of coincident and leading economic indicators. NBER Macroecon. Annu. 4, 351–394.
Sun, X., Tsang, K.P., 2017. What drives the owner-occupied and rental housing markets? Evidence from an estimated DSGE model. J. Money Credit Bank. 49 (2–3), 443–468.
Van Binsbergen, J.H., Fernández-Villaverde, J., Koijen, R.S., Rubio-Ramírez, J., 2012. The term structure of interest rates in a DSGE model with recursive preferences. J. Monetary Econ. 59 (7), 634–648.
Vázquez, J., Aguilar, P., 2021. Adaptive learning with term structure information. Eur. Econ. Rev. 134, 103689.
Walsh, C.E., 2010. Monetary Theory and Policy. The MIT Press, Cambridge, Massachusetts.
Wright, J.H., 2006. The Yield Curve and Predicting Recessions. Federal Reserve Board Finance and Economics Discussion Series Paper, FEDs Working Paper.
Wu, J.C., Xia, F.D., 2016. Measuring the macroeconomic impact of monetary policy at the zero lower bound. J. Money Credit Bank. 48 (2–3), 253–291.
Wu, J.C., Zhang, J., 2019. A shadow rate New Keynesian model. J. Econom. Dynam. Control 107, 103728.