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Monetary policy and long-term real rates ☆

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ABSTRACT

Changes in monetary policy have surprisingly strong effects on forward real rates in the distant future. A 100 basis point increase in the two-year nominal yield on a Federal Open Markets Committee announcement day is associated with a 42 basis point increase in the ten-year forward real rate. This finding is at odds with standard macro-models based on sticky nominal prices, which imply that monetary policy cannot move real rates over a horizon longer than that over which all prices in the economy can readjust. Instead, the responsiveness of long-term real rates to monetary shocks appears to reflect changes in term premia. One mechanism that could generate such variation in term premia is based on demand effects due to the existence of what we call yield-oriented investors. We find some evidence supportive of this channel.

1. Introduction

We show that changes in the stance of monetary policy have surprisingly strong effects on very distant forward real interest rates. A 100 basis point (bp) increase in the two-year nominal yield on a Federal Open Markets Committee (FOMC) announcement day, which we use as a proxy for changes in expectations regarding the path of the federal funds rate over the following several quarters, is associated with a 42 bp increase in the ten-year forward overnight real rate, extracted from the yield curve for Treasury Inflation Protected Securities (TIPS).

Our findings can be illustrated with the FOMC's announcement on January 25, 2012. On that date the FOMC significantly changed its forward guidance, indicating that it expected to hold the federal funds rate near zero “through late 2014.” It had previously stated that it expected to do so only “through mid-2013.” In response to this announcement, the expected path of short-term nominal rates fell significantly from two to five years out, with the two-year nominal yield dropping by 5 bps and the five-year nominal yield by 14 bps. More strikingly, ten-year and 20-year real forward rates declined by 5 bps and 9 bps, respectively. In other words, distant real forward rates appeared to react strongly to news about the future stance of monetary policy.

This finding is at odds with standard New Keynesian macro-models, in which the central bank's ability to influence real variables stems from that fact that goods

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prices are sticky in nominal terms. In such models, a change in monetary policy should have no impact on forward real interest rates at a horizon longer than that over which all nominal prices can readjust, and it seems implausible to think that this horizon could be anything close to ten years.[1]

So how does one make sense of our finding? One possibility is that the results are simply wrong in some sense; i.e., they are either not robust or noncausal. On the robustness front, one limitation of our analysis is that there is a relatively brief sample period in the US over which we can study real rates: TIPS were introduced in 1997, and reliable data only became available in 1999. In an effort to address this concern, we replicate our analysis on UK data over roughly the same period and find broadly similar results.

With respect to causality, a natural concern is that some of the movement in two-year nominal yields on FOMC days could be unrelated to monetary policy and could instead reflect other macro-news that also drives changes in distant forwards. If so, our estimates could suffer from an omitted variable bias. To address this concern, we can instead proxy for monetary surprises with the change in two-year nominal yields in a narrow 60-minute window surrounding FOMC announcements. When we do so, we obtain estimates that are similar to our baseline results.

Another worry is that changes in short-term nominal rates around FOMC announcements might not reflect innovations to Federal Reserve policy per se, but rather the revelation of the Fed's private information about the future evolution of the economy. For example, suppose the Fed obtains private information suggesting a permanent positive productivity shock. This shock could lead the Fed to tighten in the short run and at the same time could raise the natural (flexible-price) real interest rate in the economy forever. If so, it would be a mistake to conclude that the increase in distant forward real rates was caused by a change in monetary policy.

Although completely ruling out this possibility is difficult, we can make some progress by comparing the results we get for FOMC announcement days with the analogous results for non-FOMC days. The idea is that non-FOMC days also have their fair share of macro-news but are less likely to be informative about shifts in the Fed's reaction function. Thus, if the elasticity of long-term real rates to short-term nominal rates is simply driven by macro-news (either revealed by Fed actions or released through standard channels), this elasticity should be stronger on non-FOMC days, which arguably have a greater proportion of macro-news and less reaction-function news. However, this prediction is not borne out in the data. If anything, we find the reverse: distant forward real rates react more strongly to changes in short-term nominal rates on FOMC days than on non-FOMC days. Although not a definitive test, this finding weighs against a story based on the Fed having private information about long-run economic fundamentals.

Assuming that the results can be given a causal interpretation, what economic mechanism do they reflect? It is helpful to begin by noting that a movement in the ten-year forward real rate can always be decomposed into a change in the expected real rate that will prevail in ten years, plus a change in the ten-year real term premium. A movement in the real term premium is equivalent to saying that when the Fed raises short-term nominal rates, this increases the expected return on a carry-trade strategy that borrows short-term and buys long-term real bonds.[2]

This decomposition suggests two broad economic channels that could be at work. The first involves monetary policy somehow moving expected future real rates at very distant horizons. If this channel were operative, it would be a form of long-run monetary non-neutrality that runs directly counter to the rational-expectations spirit of New Keynesian models. In other words, it is hard to see how this channel could be squared with the bedrock assumption in these models, namely, that nominal prices are set in a rational, forward-looking manner.

The alternative possibility is that monetary policy does not move expected future real rates far out into the future but instead changes the term premia on long-term bonds. This implies that the effects on forward rates that we document should be expected to mean revert over time. To test this hypothesis, we proceed as follows. At any time t, we cumulate the changes in long-term forward rates that occurred solely on FOMC announcement days over the preceding three months. We then use these FOMC announcement day changes to forecast changes in forward rates over the subsequent 12 months. It turns out that when long-term forward rates rise on an FOMC announcement day, this predicts a reversal of forward rates over the next 12 months. The evidence is thus consistent with the proposition that monetary policy shocks induce time variation in real term premia.[3]

The question thus arises of why monetary policy could be influencing real term premia. In traditional representative-agent asset pricing models, term premia are pinned down by the covariance between real bond returns and investors' marginal utility. It is difficult to see why monetary shocks would change this covariance in the required direction, so we focus instead on an alternative class of supply and demand-based mechanisms. One specific explanation that we flesh out in detail has to do with the existence of what we call yield-oriented investors. We assume that these investors allocate their portfolios between short- and long-term bonds and, in doing so, care about current portfolio income or yield and not just expected holding-period returns. This could be because of agency or accounting considerations that lead investors to worry about short-term measures of reported performance.

A reduction in short-term nominal rates leads these investors to rebalance their portfolios toward longer-term

[1] See Clarida, Gali, and Gertler (1999) for an introduction to the New Keynesian literature and Gali (2008) for a more detailed treatment.

[2] For those more comfortable thinking in terms of stock prices, when a company's stock price goes up, one can always decompose this into news either about its expected future earnings (the analog to news about the expected future real rate here) or about its discount rate (the analog to the term premium on a carry-trade strategy).

[3] To be clear, none of our evidence directly refutes the long-run non-neutrality hypothesis that policy is somehow able to move expected real rates far out into future. Both effects could be simultaneously at work.

bonds in an effort to keep their overall portfolio yield from declining too much. This, in turn, creates buying pressure that raises the price of the long-term bonds and, hence, lowers long-term real yields and forward rates. The price pressure is independent of expectations about the actual path of future short rates; it is a pure term-premium effect. And interestingly, according to this hypothesis, conventional monetary policy moves long-term real rates in much the same way as some of the Fed's recent quantitative easing (QE) policy measures, such as its purchases of long-term Treasuries. These, too, are presumed to operate through a supply and demand effect on term premia as opposed to by changing expectations about the future path of rates.

We go on to provide some evidence that is consistent with our hypothesis about the role of yield-oriented investors. We do so by looking at the maturity of securities held by commercial banks. Banks fit with our conception of yield-oriented investors to the extent that they care about their reported earnings, which, given bank accounting rules, are based on current income from securities holdings and not mark-to-market changes in value. We find that when the yield curve steepens, banks increase the maturity of their securities holdings. Moreover, the magnitudes of these portfolio shifts are large in the aggregate, so that if they had to be absorbed by other, less yield-oriented investors (e.g., broker-dealers or hedge funds) they could plausibly drive changes in market-wide term premia. We also find that primary dealers in the Treasury market—who, unlike banks, must mark their securities holdings to market—take the other side of the trade, reducing the maturity of their Treasury holdings when the yield curve steepens.

The ideas in this paper connect to several strands of prior research. A large literature examines the impact of monetary policy surprises on long-term nominal interest rates. For example, Cochrane and Piazzesi (2002) find that a 100 bp increase in the one-month eurodollar rate around the time of a federal funds target change is associated with a 52 bp increase in ten-year nominal Treasury yields. They, too, cast this as something of a puzzle, remarking that "the size of the coefficients is particularly startling" (p. 92). In a similar vein, Gürkaynak, Sack, and Swanson (2005b) show that distant nominal forward rates respond strongly to a variety of macroeconomic news releases, including FOMC announcements.[4]

We sharpen the puzzle by focusing on real rates instead of nominal rates, which puts the long-run non-neutrality issue front and center. By contrast, Gürkaynak, Sack, and Swanson (2005b) argue that their results are consistent with a model in which long-run inflation expectations are not well anchored and are revised in light of incoming news. According to this explanation, monetary shocks could alter long-run inflation expectations but would have no impact on long-run real rates.

More recently, several papers in the monetary economics literature have also noted the surprising response of long-term real rates to monetary policy surprises. Gilchrist, Lopez-Salido, and Zakrajsek (2013) present evidence that conventional monetary policy has large effects on long-term real borrowing rates, and, like us, they argue that this occurs largely because term premia react to policy shifts. Gertler and Karadi (2013) augment a standard vector autoregression analysis of conventional monetary policy by incorporating data on the high-frequency response of interest rates to policy shocks. They find that policy shocks have a modest impact on short-term nominal rates but, nonetheless, have large effects on the real cost of long-term credit and, therefore, on real economic activity. Gertler and Karadi argue that the large response of real credit costs is due to the reaction of term premia and credit spreads, factors that are omitted from standard models of the monetary transmission mechanism.[5]

Finally, the yield-oriented investors that drive term premia in our model are reminiscent of the Rajan (2005) account of investor behavior in a low interest rate environment. And the idea that supply and demand effects can have important consequences in the Treasury market is central to a number of recent papers, including Vayanos and Vila (2009), Greenwood and Vayanos (2010, 2014), Krishnamurthy and Vissing-Jorgensen (2011, 2012), Gagnon, Raskin, Remache, and Sack (2011), and Hanson (2014). An important antecedent to this work is Modigliani and Sutch (1966).

The remainder of the paper is organized as follows. In Section 2, we document the strong sensitivity of long-term real forward rates to monetary policy news and argue that this relation is likely to be causal. In Section 3, we make the case that movements in long-term forward rates around monetary policy announcements reflect changes in term premia. In Section 4, we investigate the mechanism behind these changing term premia. Section 5 concludes.

2. The sensitivity of long-term real forward rates to monetary policy news

We begin by documenting the surprising sensitivity of distant real forward rates to monetary policy shocks. We then argue that this relation is likely to be causal.

2.1. Measuring monetary policy news

To get started, we need a measure of monetary policy news. A growing consensus exists that changes in the policy outlook are the primary form of monetary policy news on FOMC announcement days. Thus, building on Gürkaynak, Sack, and Swanson (2005a) and Campbell, Evans, Fisher, and Justiniano (2012), our measurement strategy is based on the premise that, at least since 1994, a significant portion of the news contained in FOMC announcements is about the expected path of the federal

[4] Other papers in this tradition include Cook and Hahn (1989), Evans and Marshall (1998), and Kuttner (2001).

[5] Instead of reflecting changes in term premia, Nakamura and Steinsson (2013) argue that the large response of distant real forwards to policy surprises reflects the fact that nominal price rigidities are far more severe than typically assumed. This implies that monetary policy is not neutral even at fairly long horizons.

funds rates over the next several quarters as opposed to surprise changes in the current federal funds rate.[6]

To capture revisions to the full expected path of the funds rate over the coming quarters in a simple and transparent manner, we use the change in the two-year nominal Treasury yield on FOMC announcement dates as our proxy for monetary policy news. However, as described in our robustness tests below, we obtain similar results with a variety of related variables that capture revisions in expected short rates over the following several quarters. The key is that these variables capture news about the expected medium-term path of interest rates as opposed to news only about rates over the coming month or two.

We use data from Gürkaynak, Sack, and Wright (2007, 2010) on the nominal Treasury yield curve and the real (TIPS) Treasury yield curve as updated regularly by the Federal Reserve Board. For each day, they estimate the six-parameter model of the instantaneous forward curve proposed by Svensson (1994). Zero-coupon yields are then obtained by integrating along the estimated forward curve

$$y_t^{(n)} = n^{-1} \int_0^n f_t^{(m)}\, dm. \tag{1}$$

We can decompose the n-year nominal forward rate $f_t^{\$(n)}$ into the sum of the forward real rate $f_t^{TIPS(n)}$ and the forward break-even inflation rate $f_t^{\pi(n)}$,

$$f_t^{\$(n)} = f_t^{TIPS(n)} + f_t^{\pi(n)}. \tag{2}$$

The n-year nominal zero-coupon yield can be decomposed analogously:

$$y_t^{\$(n)} = y_t^{TIPS(n)} + y_t^{\pi(n)}. \tag{3}$$

In our baseline specification, for an FOMC meeting on day t, we compute changes from t–1 to $t+1$ to capture the full market response to the announcement. Our implicit assumption is that the full reaction to an FOMC announcement might not be instantaneous, particularly for long-term yields. This could be because investors are uncertain about the implications of a given piece of news and update their beliefs as others' interpretations are revealed via trading volume, the price process, and the financial media. Thus, it could take some time for the market to digest the information content of an announcement.

The Treasury market microstructure literature is consistent with this view. Fleming and Remolona (1999) find that price formation is gradual with heightened levels of volume and volatility lasting 90 or more minutes following major announcements. More relevant for us, Gürkaynak, Sack, and Swanson (2005a) find that it takes markets time to impound news about the future path of rates contained in FOMC statements, but it takes almost no time to impound news about the current target. Said differently, it appears to take longer-term yields more time to fully react to FOMC announcements.

Given this evidence, we want to choose a window long enough to span the period of elevated post-announcement price volatility. In this context, the timing of our daily Treasury data argues in favor of using a two-day window. Most FOMC announcements in our sample are at 2:15 p.m., and the Treasury quotes underlying our fitted yields curves are taken from 3:00 p.m. closing prices. As a result, a one-day horizon would allow only 45 min for long-term yields to adjust. Our results our qualitatively similar but somewhat smaller in magnitude, if we instead measure changes over the one-day interval from day t–1 to t.

2.2. Baseline results for the US

In our baseline specifications, we regress changes in forward nominal rates, forward real rates, and forward break-even inflation rates on changes in two-year nominal yields

$$\Delta f_t^{\$(n)} = a_{\$}(n) + b_{\$}(n)\Delta y_t^{\$(2)} + \Delta\varepsilon_t^{\$(n)} \tag{4}$$

$$\Delta f_t^{TIPS(n)} = a_{TIPS}(n) + b_{TIPS}(n)\Delta y_t^{\$(2)} + \Delta\varepsilon_t^{TIPS(n)}, \tag{5}$$

and

$$\Delta f_t^{\pi(n)} = a_{\pi}(n) + b_{\pi}(n)\Delta y_t^{\$(2)} + \Delta\varepsilon_t^{\pi(n)}. \tag{6}$$

We focus on FOMC announcement dates from 1999 to February 2012. We exclude five FOMC announcement dates that contained significant news about the Fed's large-scale asset purchases (LSAPs; sometimes referred to as QE1, QE2 and Operation Twist).[7] We do so because the mechanism underlying long-term rate movements on these dates is potentially different from that driving market reactions to more conventional FOMC announcements.

Table 1 and Fig. 1 present the basic results. Panel A of Fig. 1 shows how the nominal forward curve responds to a 100 bp shock to short-term nominal rates. It plots the coefficients from Eq. (4) for maturities $n=5,\ldots,20$ along with 95% confidence intervals. Panel B of Fig. 1 decomposes the response of nominal forwards into a change in real forwards and forward break-even inflation, plotting the coefficients from Eqs. (5) and (6). By construction, the sum of the two coefficients shown in Panel B equals the coefficient in Panel A. Table 1 lists all the regression coefficients.

Table 1 and Fig. 1 show that distant nominal forwards respond significantly to changes in short-term nominal rates on FOMC days. And, surprisingly, this response is driven almost exclusively by movements in real forwards. A 100 bp shock to the two-year nominal rate on an FOMC announcement date is associated with a 45 bp increase in ten-year nominal forwards ($t=3.54$). And this 45 bp increase can be decomposed into a 42 bp rise in real

[6] In 1994, the FOMC began issuing a press release with the current federal funds target after every meeting and also began releasing announcements discussing the economic and policy outlook. Prior to 1994, the FOMC implicitly announced the change in its target via the size and type of the next open-market operation following a policy change (typically the day after the FOMC meeting). From 1994 to mid-1999, the FOMC released a statement only when it changed the policy target. However, since mid-1999, the FOMC has released a statement following each meeting.

[7] The five excluded FOMC announcement dates are March 18. 2009 (QE1), August 10, 2010 (QE2), September 21, 2010 (QE2), November 3, 2010 (QE2), and September 21, 2011 (Operation Twist). Our results are robust both to including these dates and to excluding others (December 16, 2008 and January 28, 2009) that arguably also contained some information about the LSAPs.

Table 1
Response of US Treasury forward rates to monetary policy news.
Regressions of changes in nominal, real, and break-even inflation instantaneous forward rates ($X=\$$, *TIPS*, and π) on changes in the two-year nominal yield on Federal Open Markets Committee (FOMC) announcement days from 1999 through February 2012:

$$\Delta f_t^{X(n)} = a_X(n) + b_X(n)\Delta y_t^{\$(2)} + \Delta\varepsilon_t^{X(n)}.$$

We estimate these regressions for maturities of $n=5,\ldots,20$. For an announcement on day t, we compute the two-day change from $t-1$ to $t+1$. t-Statistics, based on robust standard errors, are shown in brackets. We exclude five FOMC announcements dates from 2009 to 2011, when there was significant news about the Federal Reserve's large scale asset purchase (LSAP) programs. Daily estimates of nominal forward rates, real forward rates, and break-even inflation forward rates are based on Gürkaynak, Sack, and Wright (2007, 2010). The data, updated regularly by Federal Reserve Board staff, are available at http://www.federalreserve.gov/pubs/feds/2006/200628/200628abs.html and http://www.federalreserve.gov/pubs/feds/2008/200805/200805abs.html.

	Nominal forwards			Real forwards			Inflation forwards		
n	$b_\$(n)$	$[t]$	R^2	$b_{TIPS}(n)$	$[t]$	R^2	$b_\pi(n)$	$[t]$	R^2
5	0.843	[6.07]	0.30	0.653	[5.98]	0.24	0.190	[2.46]	0.05
6	0.729	[4.90]	0.21	0.563	[5.77]	0.20	0.166	[1.80]	0.04
7	0.634	[4.22]	0.16	0.505	[5.58]	0.18	0.129	[1.22]	0.02
8	0.557	[3.84]	0.13	0.467	[5.24]	0.17	0.090	[0.80]	0.01
9	0.496	[3.64]	0.11	0.441	[4.88]	0.17	0.055	[0.48]	0.00
10	0.446	[3.54]	0.09	0.421	[4.63]	0.18	0.025	[0.23]	0.00
11	0.405	[3.47]	0.09	0.405	[4.47]	0.18	0.001	[0.01]	0.00
12	0.371	[3.37]	0.08	0.390	[4.37]	0.18	−0.018	[−0.18]	0.00
13	0.342	[3.21]	0.07	0.376	[4.29]	0.17	−0.034	[−0.35]	0.00
14	0.315	[2.99]	0.07	0.362	[4.21]	0.17	−0.047	[−0.48]	0.00
15	0.291	[2.73]	0.06	0.350	[4.10]	0.15	−0.059	[−0.58]	0.00
16	0.267	[2.45]	0.06	0.338	[3.95]	0.14	−0.071	[−0.65]	0.01
17	0.244	[2.17]	0.05	0.327	[3.76]	0.13	−0.083	[−0.70]	0.01
18	0.222	[1.89]	0.04	0.316	[3.56]	0.12	−0.094	[−0.74]	0.01
19	0.199	[1.62]	0.04	0.306	[3.36]	0.11	−0.107	[−0.77]	0.01
20	0.176	[1.36]	0.03	0.296	[3.15]	0.09	−0.120	[−0.79]	0.01

forwards ($t=4.63$) and a 3 bp rise in forward break-even inflation ($t=0.23$). This pattern holds even as we consider more distant forwards. A 100 bp shock to two-year nominal rates is associated with an 18 bp increase in 20-year nominal forwards ($t=1.32$), which reflects a 30 bp rise in real forwards ($t=3.15$) and a 12 bp decline in forward break-even inflation ($t=-0.79$).

Table 2 conducts a variety of robustness exercises. First, we vary the event window. In our baseline results, we use a two-day window from day $t-1$ to day $t+1$. We also report comparable results when using a one-day window from t–1 to t. This leads to somewhat smaller effects: a 100 bp shock to the two-year nominal rate is now associated with a 25 bp rise in ten-year nominal forwards ($t=2.98$), a 22 bp rise in real forwards ($t=2.90$), and a 2 bp rise in forward break-even inflation ($t=0.30$).[8]

Next we try using a variety of alternative measures of short-run nominal rates in place of two-year Treasury yields: one-year Treasury yields, federal funds futures, and eurodollar futures contracts. The basic take-away is that everything works similarly with any variable that captures news about the medium-term path of rates as opposed to one that captures only what the target will be in the next few weeks. This is shown explicitly when we construct the “future path of policy” news factor as in Gürkaynak, Sack, and Swanson (2005a). We also vary the sample. For example, we add dates with major news about QE policies or the dates on which FOMC minutes are released. This has little impact on the results.

Finally, because TIPS are known to carry a smaller liquidity premium than nominal Treasuries, one could wonder whether we obtain a similar decomposition of nominal forwards into real and break-even inflation using inflation swaps as opposed to TIPS.[9] A persistent liquidity differential is not a concern given our high-frequency empirical strategy. However, we want to know if our results reflect monetary policy-induced changes in liquidity premia. We investigate this in two ways. First, we use inflation swap yields and nominal yields to back out a synthetic real yield. Second, we examine whether proxies for the equilibrium price of liquidity do, in fact, respond to monetary policy.

The last row in Table 2 uses data on zero-coupon inflation swaps to construct a synthetic real forward, defined as the forward rate implied by nominal Treasuries less forward inflation implied by swaps. This approach yields point estimates that are somewhat larger than those based on TIPS, though the difference is not statistically significant.

We examine the behavior of a standard proxy for the price of liquidity: the yield spread between off-the-run and on-the-run Treasuries (Krishnamurthy, 2002). We regress

[8] The decline in the coefficient is largely due to the use of a two-day window for long-term yields on the left-hand side of the regression. If we use a two-day window for long-term yields on the left-hand side and a one-day window for short-term yields on the right-hand side, we obtain $b=0.414$ ($t=3.04$), which is very close to our baseline result.

[9] TIPS are very liquid, but nominal Treasuries are the most liquid asset class in global markets. As a result, nominal Treasuries command a liquidity premium relative to private bonds (Krishnamurthy and Vissing-Jorgensen, 2012) as well as relative to TIPS (Fleckenstein, Longstaff, and Lustig, 2014; Pflueger and Viceira, 2013).

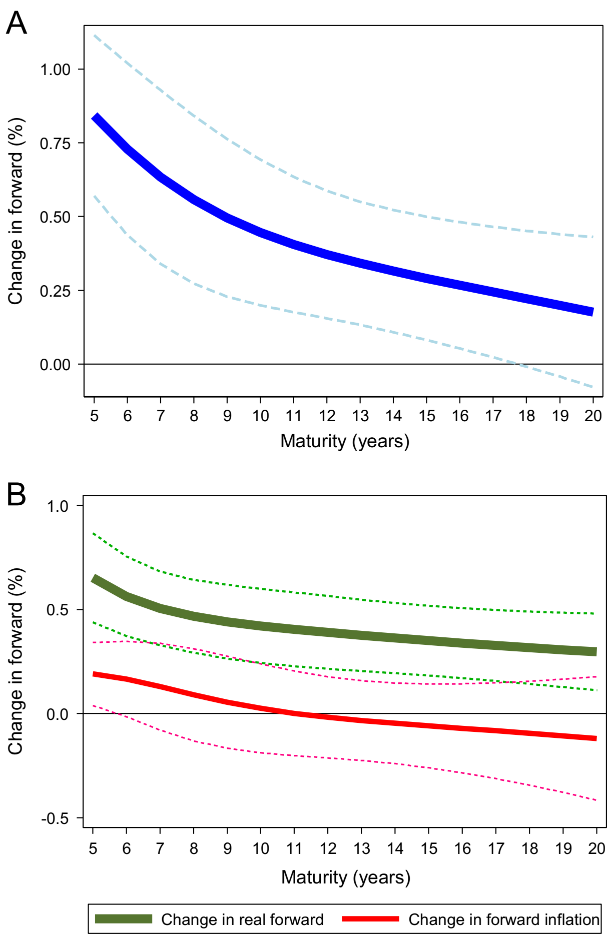

Fig. 1. Response of US forwards to monetary policy news on Federal Open Markets Committee (FOMC) days. Panel A plots the coefficient $b_{\$}(n)$ from estimating Eq. (4) on FOMC announcement dates from 1999 to February 2012:

$$\Delta f_t^{\$(n)} = a_{\$}(n) + b_{\$}(n)\Delta y_t^{\$(2)} + \Delta\varepsilon_t^{\$(n)}.$$

Panel B plots the coefficients $b_{TIPS}(n)$ and $b_{\pi}(n)$ from estimating Eqs. (5) and (6) on FOMC announcement dates from 1999 to February 2012:

$$\Delta f_t^{TIPS(n)} = a_{TIPS}(n) + b_{TIPS}(n)\Delta y_t^{\$(2)} + \Delta\varepsilon_t^{TIPS(n)} \text{ and } \Delta f_t^{\pi(n)} = a_{\pi}(n) + b_{\pi}(n)\Delta y_t^{\$(2)}{}_t + \Delta\varepsilon_t^{\pi(n)}.$$

Confidence intervals, based on robust standard errors, are shown as dashed lines. (A) Response of nominal forwards by maturity. (B) Response of real and break-even inflation forwards by maturity.

the change in the yield spread between the old on-the-run and current on-the-run ten-year nominal Treasury on the change in two-year nominal yields around FOMC announcements. Doing so, we find little evidence that monetary surprises impact the price of liquidity: the estimated coefficient is $b = -0.004$ ($t = -0.39$). In combination, these exercises suggest that changes in liquidity premia play little role in explaining our results.

2.3. Parallel results for the UK

To further investigate the robustness of our results, we run the analogous set of regressions using UK data. To do so, we rely on the yield curve estimates published by the Bank of England (BOE), which employ the spline-based techniques described in Anderson and Sleath (1999). We estimate Eqs. (4)–(6) on all monetary policy announcement dates since 1994.[10] Our proxy for news on announcement dates is the change in the two-year nominal yield. We compute changes from $t-1$ to $t+1$ for meetings on day t. And we drop six announcement dates from 2009 to 2011, when there was significant news about the BOE's quantitative easing operations.[11]

Table 3 and Fig. 2 present the basic results for the UK. The estimates are qualitatively similar to those from the US, although the magnitude of the effect is somewhat

[10] Although the UK has issued inflation-linked bonds since 1985, UK authorities began holding regularly scheduled monetary policy meetings analogous to those held by the FOMC only in 1994, so we begin our analysis then.

[11] The list is based on Table A in Joyce, Tong, and Woods (2011). The dates are March 5, 2009, May 7, 2009, August 6, 2009, November 5, 2009, February 4, 2010, and October 6, 2011.

Table 2
Robustness checks for US.

Regressions of changes in nominal, real, and break-even inflation instantaneous forward rates (X=\$, *TIPS*, and π) on changes in various short rates on Federal Open Markets Committee (FOMC) announcement days. t-Statistics, based on robust standard errors, are shown in brackets. We first vary the window (one-day versus two-day changes) used to compute changes in long-term forwards and short-term rates. We next use a variety of different proxies for monetary policy news on FOMC announcement dates, including the "future path of policy" news factor as in Gürkaynak, Sack, and Swanson (2005a). Data on federal funds futures and eurodollar futures are from Bloomberg. Next, we vary the sample. Finally, we address concerns about liquidity effects using data on inflation swaps. We work with ten-year rates ten-year forward here. The Treasury Inflation Protected Securities (TIPS)-implied forward real rates are from Gürkaynak, Sack, and Wright (2010). We use data on zero-coupon inflation swaps from Bloomberg to construct proxies for real forwards and forward inflation that do not rely on TIPS data. Our proxy for the real forward rate is the difference between nominal Treasury forwards from Gürkaynak, Sack, and Wright (2007) and forward inflation computed using zero-coupon inflation swaps. Because inflation swap data are available only beginning in July 2004 and are sporadic until August 2005, these regressions use 57 observations.

		Nominal forwards			Real forwards			Inflation forwards		
	Observations	$b_{\$}(10)$	$[t]$	R^2	$b_{TIPS}(10)$	$[t]$	R^2	$b_{\pi}(10)$	$[t]$	R^2
Baseline:										
Two-year UST	107	0.446	[3.54]	0.09	0.421	[4.63]	0.18	0.025	[0.23]	0.00
Vary proxy for Monetary Policy news										
Two-year US Treasury, one-day change	107	0.245	[2.98]	0.08	0.215	[2.90]	0.13	0.021	[0.30]	0.00
One-year US Treasury	107	0.186	[1.30]	0.01	0.287	[2.60]	0.07	−0.100	[−0.89]	0.01
One-year US Treasury, one-year forward	107	0.505	[5.13]	0.17	0.408	[5.42]	0.24	0.097	[1.14]	0.02
Three-quarter forward eurodollar	107	0.239	[2.39]	0.07	0.286	[4.43]	0.20	−0.046	[−0.66]	0.01
Six-quarter forward eurodollar	107	0.398	[4.23]	0.23	0.326	[6.33]	0.33	0.072	[0.95]	0.02
Nine-month forward fed funds	89	0.258	[1.85]	0.03	0.354	[3.72]	0.11	−0.096	[−0.82]	0.01
12-month forward fed funds	74	0.308	[1.90]	0.04	0.401	[3.42]	0.14	−0.093	[−0.75]	0.01
Gürkaynak, Sack, and Swanson (2005a) path factor	107	0.274	[2.43]	0.10	0.279	[3.75]	0.23	−0.005	[−0.07]	0.00
Vary sample										
Add five quantitative easing dates	112	0.504	[3.30]	0.10	0.455	[4.97]	0.18	0.049	[0.40]	0.00
Regular meeting	100	0.452	[4.16]	0.09	0.383	[4.04]	0.14	0.069	[0.73]	0.01
Unscheduled meeting	7	0.565	[2.62]	0.38	0.668	[3.66]	0.67	−0.103	[−0.38]	0.02
Add minutes dates	212	0.528	[5.85]	0.17	0.312	[4.76]	0.14	0.203	[2.04]	0.04
Explore liquidity effects										
TIPS-implied forwards	57	0.255	[1.43]	0.04	0.440	[2.97]	0.16	−0.186	[−1.11]	0.04
Inflation swap-implied forwards	57	0.255	[1.43]	0.04	0.838	[2.41]	0.20	−0.583	[−1.50]	0.12

Table 3
Response of UK gilt forward rates to monetary policy news.

Regressions of changes in nominal, real, and break-even inflation instantaneous forward rates (X=\$, *TIPS*, and π) on changes in the two-year nominal gilt yield on UK monetary policy announcement days from 1994 to February 2012:

$$\Delta f_t^{X(n)} = a_X(n) + b_X(n)\Delta y_t^{\$(2)} + \Delta\varepsilon_t^{X(n)}.$$

We estimate these regressions for maturities of n=5,…,20. For an announcement on day t, we compute the two-day change from $t-1$ to $t+1$. t-Statistics, based on robust standard errors, are shown in brackets. Beginning in June 1997, our policy announcement dates correspond to meetings of the Bank of England (BOE) Monetary Policy Committee (MPC) available at http://www.bankofengland.co.uk/monetarypolicy/Pages/decisions.aspx. From January 1994 to May 1997, we use the dates of the Monthly Monetary Meetings between the Governor of the BOE and the Chancellor of the Exchequer from Table 6.1 of Cobham (2002). We exclude six MPC announcements dates when there was significant news about the BOE's quantitative easing operations. The UK yield curve data are based on the methods described in Anderson and Sleath (1999). The data are available at http://www.bankofengland.co.uk/statistics/Pages/yieldcurve/default.aspx.

	Nominal forwards			Real forwards			Inflation forwards		
n	$b_{\$}(n)$	$[t]$	R^2	$b_{TIPS}(n)$	$[t]$	R^2	$b_{\pi}(n)$	$[t]$	R^2
5	0.591	[7.24]	0.25	0.396	[8.04]	0.30	0.196	[3.07]	0.05
6	0.510	[6.44]	0.18	0.356	[7.85]	0.27	0.154	[2.54]	0.03
7	0.445	[5.74]	0.13	0.324	[7.88]	0.25	0.121	[1.97]	0.02
8	0.392	[5.15]	0.11	0.298	[7.81]	0.23	0.094	[1.49]	0.01
9	0.345	[4.61]	0.09	0.274	[7.48]	0.21	0.071	[1.08]	0.01
10	0.300	[4.05]	0.07	0.254	[6.89]	0.19	0.046	[0.69]	0.00
11	0.258	[3.43]	0.05	0.236	[6.16]	0.17	0.021	[0.30]	0.00
12	0.218	[2.80]	0.04	0.221	[5.45]	0.14	−0.003	[−0.04]	0.00
13	0.181	[2.23]	0.03	0.207	[4.83]	0.13	−0.026	[−0.31]	0.00
14	0.151	[1.76]	0.02	0.195	[4.33]	0.11	−0.045	[−0.50]	0.00
15	0.126	[1.41]	0.01	0.185	[3.93]	0.10	−0.059	[−0.62]	0.00
16	0.108	[1.17]	0.01	0.177	[3.63]	0.09	−0.069	[−0.70]	0.01
17	0.095	[1.01]	0.01	0.171	[3.40]	0.08	−0.076	[−0.73]	0.01
18	0.088	[0.92]	0.01	0.166	[3.23]	0.07	−0.078	[−0.74]	0.01
19	0.085	[0.89]	0.01	0.162	[3.10]	0.07	−0.077	[−0.72]	0.01
20	0.085	[0.89]	0.01	0.165	[3.08]	0.07	−0.077	[−0.70]	0.01

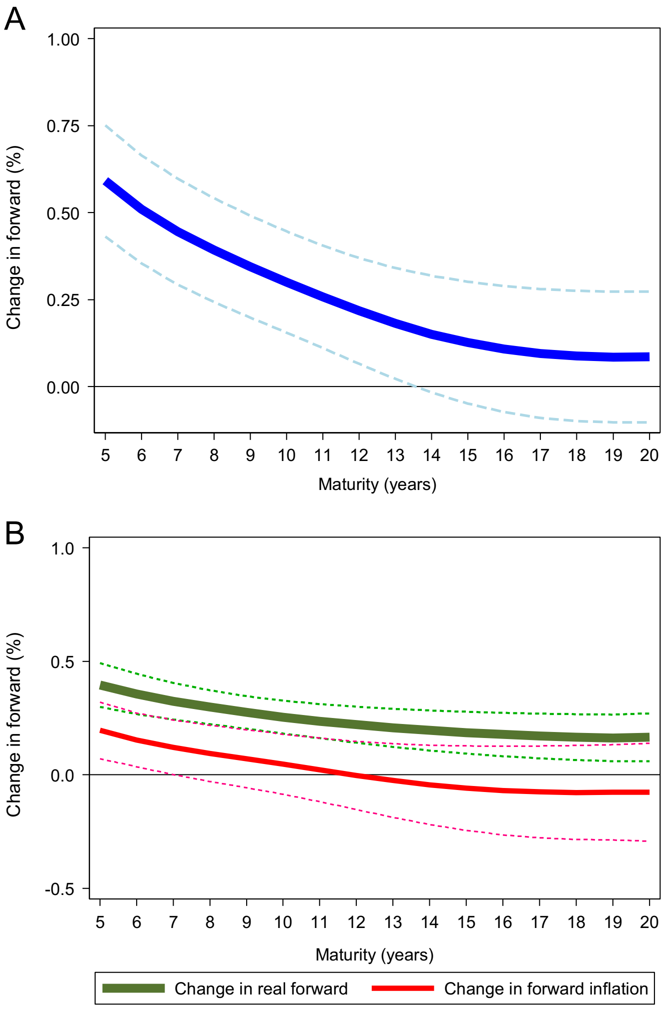

Fig. 2. Response of UK forwards to monetary policy news on announcement days. Panel A plots the coefficient $b_{\$}(n)$ from estimating Eq. (4) on UK monetary policy announcement dates from 1994 to February 2012:

$$\Delta f_t^{\$(n)} = a_{\$}(n) + b_{\$}(n)\Delta y_t^{\$(2)} + \Delta\varepsilon_t^{\$(n)}.$$

Panel B plots the coefficients $b_{TIPS}(n)$ and $b_{\pi}(n)$ from estimating Eqs. (5) and (6) on UK monetary policy announcement dates from 1994 to February 2012:

$$\Delta f_t^{TIPS(n)} = a_{TIPS}(n) + b_{TIPS}(n)\Delta y_t^{\$(2)} + \Delta\varepsilon_t^{TIPS(n)} \text{ and } \Delta f_t^{\pi(n)} = a_{\pi}(n) + b_{\pi}(n)\Delta y_t^{\$(2)} + \Delta\varepsilon_t^{\pi(n)}.$$

Confidence intervals, based on robust standard errors, are shown as dashed lines. (A) Response of nominal forwards by maturity. (B) Response of real and break-even inflation forwards by maturity.

smaller in the UK. In particular, for the ten-year forward real rate, the coefficient on the two-year nominal yield is 0.254 in the UK as compared with 0.421 in the US.

2.4. Do monetary policy shocks cause the movements in distant real forward rates?

One could worry that some of the movements in two-year yields on FOMC days are due not to monetary policy surprises but rather other fundamental macro-news that also impacts distant forwards. Because we do not control for other macro-news, our ordinary least squares (OLS) regressions will yield biased estimates of the effect of monetary policy on distant real forwards if fundamental macro-news has a different effect on forwards than monetary policy. To deal with this concern, we follow Gertler and Karadi (2013) and Gilchrist, Lopez-Salido, and Zakrajsek (2013) and estimate our baseline specifications using the intraday change in two-year yields in a narrow 60-minute window around each FOMC announcement as an instrument for the two-day change in two-year yields.[12] The exclusion restriction here

[12] We obtain the precise announcement times from Gürkaynak, Sack, and Swanson (2005a) and Lucca and Moench (2013). Given the microstructure evidence, we use the 60-minute announcement window from Gürkaynak, Sack, and Swanson (2005a), which begins 15 min prior to the announcement and ends 45 min after. We are grateful to Refet Gürkaynak for sharing his intraday data on yield changes surrounding FOMC announcements. The underlying data source for intraday changes in Treasury yields is GovPX.

Table 4
Instrumental variables (IV) estimates of the response of ten-year US Treasury forward rates to monetary policy news.

This table shows regressions of changes in ten-year nominal, real, and break-even inflation forward rates (X=\$, *TIPS*, and π) on changes in the two-year nominal yield from 1999 through February 2012:

$$\Delta f_t^{X(10)} = a_X(10) + b_X(10)\Delta y_t^{\$(2)} + \Delta\varepsilon_t^{X(10)}.$$

Row 1 shows ordinary least squares (OLS) estimates that restrict attention to Federal Open Markets Committee (FOMC) announcement dates. Rows 2 and 3 show the corresponding IV estimates. In Row 2, we instrument for the change in the two-year yield using the change in the two-year yields in a narrow 60-minute window surrounding the announcements (60-m Δ). Row 3 uses the 60-minute change as well as the 60-minute change raised to the second and third powers (60-m Δ poly). *t*-Statistics, based on robust standard errors, are shown in brackets. We also report the first stage *F*-statistic on the instruments. We do not have intraday data for three FOMC announcements in our sample, so all estimates in this table are based on 104 announcement dates.

Specification			Nominal forwards			Real forwards			Inflation forwards		
Estimator	Instruments	1st Stage *F*	$b_\$(10)$	$[t]$	R^2	$b_{TIPS}(10)$	$[t]$	R^2	$b_\pi(10)$	$[t]$	R^2
OLS	–	–	0.415	[3.32]	0.09	0.422	[4.58]	0.18	−0.007	[−0.06]	0.00
IV	60-m Δ	20.64	0.381	[1.05]	0.09	0.561	[2.15]	0.16	−0.180	[−0.98]	0.00
IV	60-m Δ poly	13.11	0.465	[1.45]	0.08	0.612	[2.76]	0.14	−0.147	[−0.83]	0.00

is that movements in two-year yields in this 60-minute window solely reflect monetary policy surprises. This seems plausible because almost all FOMC announcements in our sample occur at roughly 2:15 p.m., macroeconomic data is almost always released at 8:30 a.m. or 10:00 a.m., and almost all major corporate news is released after stock exchanges close at 4:00 p.m.

As shown in Row 2 of Table 4, this instrumental variables (IV) procedure produces point estimates that are a bit larger than our baseline OLS estimates. Following Gilchrist, Lopez-Salido, and Zakrajsek (2013), we add squares and cubes of the intraday change as instruments in Row 3 because they add explanatory power for the two-day change in two-year yields. Using these additional instruments has little effect on our IV estimates. Fig. 3 redoes Fig. 1 with this instrumental variables estimator. In summary, our results are similar whether we measure monetary policy surprises using two-day changes or using 60-minute intraday changes. In this sense, our findings are consistent with those of Gürkaynak, Sack, and Swanson (2005a), who, after comparing daily and intraday data, conclude that "the surprise component of monetary policy announcements can be measured very well using just daily data" (p. 66).

A distinct concern is that the Fed's policy announcement is simply a response to its private information about the future evolution of the economy, and it is the release of the Fed's private information—as opposed to news about its reaction function—that moves long-term real rates. For example, suppose the Fed has private information that the economy's long-run growth potential is weaker than previously believed. This could cause the Fed to ease policy, reducing the expected path of nominal rates over the next several quarters. And, once disclosed, the same information could also lead investors to expect the long-run natural real rate to decline. However, the movement in long-term real rates would not be a causal consequence of monetary policy in this case, as it would have happened even had the Fed chosen not to ease.

This reverse-causality story is already somewhat suspect on an a priori basis, because it presumes that the Fed has material private information about the very long-run evolution of the economy. And a variety of studies have shown that the Fed does not have any forecasting advantage relative to private analysts more than a few quarters into the future.[13]

Nevertheless, we take a crude stab at testing this reverse-causality hypothesis. To do so, we compare our results with those on all non-FOMC announcement days. The intuition for this experiment is as follows. Non-FOMC days see the release of a variety of fundamental macro-news items (the same kind of macro-news that the Fed is ostensibly revealing with its FOMC announcements in the private-information story) but are less likely to bring news about the Fed's reaction function. Thus, if the elasticity of long-term real rates to short-term nominal rates is simply driven by macro-news, as is posited in the reverse-causality hypothesis, this elasticity should be stronger on non-FOMC days, which arguably have a greater relative proportion of macro-news as compared with reaction-function news.

To implement the test, we estimate

$$\Delta f_t^{TIPS(n)} = a + b\Delta y_t^{\$(2)} + c\,FOMC_t + d\Delta y_t^{\$(2)}\,FOMC_t + \varepsilon_t^{TIPS(n)}, \tag{7}$$

for n=5, 10, and 20, using all days in the sample. The results are displayed in Table 5. The key coefficient of interest is that on the interaction term, d, which captures how the elasticity of long-term real forward rates to short-term nominal rates on FOMC days differs from that on non-FOMC days. According to the reverse-causality hypothesis, this coefficient should be negative. In fact, it is generally positive, although only marginally significant.

[13] Romer and Romer (2000) argue that Fed inflation forecasts for the coming quarters outperformed those of private forecasters from the late 1960s to the early 1990s. By contrast, Faust, Swanson, and Wright (2004) argue that FOMC policy surprises contain little information that could be used to improve macroeconomic forecasts and that private forecasters do not appear to revise their forecasts in response to policy surprises. Regardless, no argument appears in the literature that the Fed has a significant forecasting advantage at anything close to a ten-year horizon.

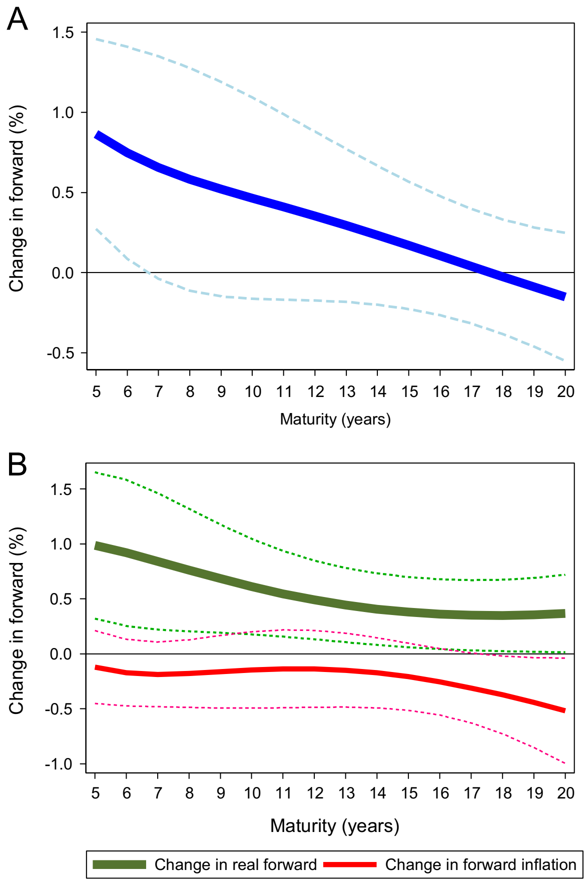

Fig. 3. Instrumental variables (IV) estimates of the response of US Treasury forward rates to monetary policy news. Panel A plots the coefficient $b_{\$}(n)$ from IV estimation of Eq. (4) on Federal Open Markets Committee announcement dates from 1999 to February 2012:

$$\Delta f_t^{\$(n)} = a_{\$}(n) + b_{\$}(n)\Delta y_t^{\$(2)} + \Delta\varepsilon_t^{\$(n)}.$$

Panel B plots the coefficients $b_{TIPS}(n)$ and $b_{\pi}(n)$ from IV estimation of Eqs. (5) and (6):

$$\Delta f_t^{TIPS(n)} = a_{TIPS}(n) + b_{TIPS}(n)\Delta y_t^{\$(2)} + \Delta\varepsilon_t^{TIPS(n)} \text{ and } \Delta f_t^{\pi(n)} = a_{\pi}(n) + b_{\pi}(n)\Delta y_t^{\$(2)}{}_t + \Delta\varepsilon_t^{\pi(n)}.$$

For an announcement on day t, we compute the two-day change from $t-1$ to $t+1$. We instrument for the two-day change in the two-year yield using the change in narrow 60-minute window surrounding the announcement as well as the 60-minute change raised to the second and third powers. t-Statistics, based on robust standard errors, are shown in brackets. (A) Response of nominal forwards by maturity. (B) Response of real and break-even inflation forwards by maturity.

The point estimates for ten-year real forwards suggest that the elasticity on non-FOMC days is 0.268 as compared with a value of 0.421 on FOMC days.

Thus, the results in Table 5 fail to support the reverse-causality hypothesis. However, this is not the same thing as having a clean instrument for exogenous shocks to the Fed's reaction function. So, while we believe the balance of the evidence favors a causal interpretation of the role of monetary policy on long-term real forwards, the identification is admittedly not airtight.

Table 5
Response of US long-term forward rates to changes in short-term rates.

Regressions of changes real instantaneous forward rates on changes in short-term nominal rates on all days, allowing for a differential response on Federal Open Markets Committee (FOMC) announcement dates (excluding any quantitative easing dates):

$$\Delta f_t^{TIPS(n)} = a + b\Delta y_t^{\$(2)}{}_t + c\,FOMC_t + d\Delta y_t^{\$(2)}\,FOMC_t + \varepsilon_t^{X(n)}.$$

We estimate these regressions for five-, ten-, and 20-year forwards using daily data from 1999 through February 2012. Standard errors are based on Newey and West (1987) standard errors allowing for serial correlation at up to two lags.

	Five-year forwards	Ten-year forwards	20-year forwards
$\Delta y^{\$(2)}$	0.493	0.268	0.240
	[15.90]	[12.30]	[9.32]
FOMC	−0.003	−0.005	−0.010
	[−0.26]	[−0.58]	[−1.20]
$\Delta y^{\$(2)}$ *FOMC*	0.160	0.153	0.057
	[1.47]	[1.69]	[0.60]
Constant	−0.001	−0.001	−0.000
	[−0.61]	[−0.38]	[−0.23]
Number of observations	3,283	3,283	3,283
R^2	0.22	0.11	0.05

3. Changes in expected future rates versus changes in term premia

If one accepts the premise that monetary policy has an important causal impact on long-term real forward rates, then the natural question to ask is whether this reflects changes in expected future real rates or changes in term premia. If it is the former, this would represent a direct challenge to the notion that monetary policy is neutral in the long run, because the implication would be that a change in policy today has a large effect on the expected level of the real rate ten years or more into the future. If it is the latter, this opens the door to a novel monetary transmission channel. And one would then want to understand the strength and persistence of this term premium effect as well as the economic mechanisms that give rise to it.

As a matter of bond accounting, a change in the n-year forward rate can always be decomposed into a change in the expected rate that will prevail in n–1 years plus a change in the n-year term premium.[14] Letting $f_t^{(n)}$ be the n-year forward rate at time t, $r_{t+1}^{(n)}$ the realized return on an n-period zero-coupon bond from t to $t+1$, and $y_t^{(1)}$ the yield on a one-period bond at time t, it is easy to show

[14] We work with one-year forward rates here as opposed to the instantaneous forward rates used above. We do this to exploit the simple decompositions for one-year forward rates, but this has only a trivial impact on the estimates. Formally, we have $f_t^{(n)} \equiv ny_t^{(n)} - (n-1)y_t^{(n-1)}$ and $r_{t+1}^{(n)} \equiv ny_t^{(n)} - (n-1)y_{t+1}^{(n-1)}$. Adding and subtracting terms yields $f_t^{(n)} = f_{t+1}^{(n-1)} + [r_{t+1}^{(n)} - r_{t+1}^{(n-1)}]$. Iterating forward implies $f_t^{(n)} = y_{t+n-1}^{(1)} + \sum_{j=1}^{n-1} [r_{t+j}^{(n+1-j)} - r_{t+j}^{(n-j)}]$. Note, too, that Eq. (8) is strictly true only over short intervals in which expected excess returns are near zero. More generally, only unexpected changes in forwards—equivalently, unexpected bond returns—contain news.

that, for changes in distant forward rates over a short horizon,

$$\Delta f_t^{(n)} = \overbrace{\Delta E_t[y_{t+n-1}^{(1)}]}^{\text{News about future short rates}} + \overbrace{\Delta E_t[\textstyle\sum_{j=1}^{n-1}(r_{t+j}^{(n+1-j)} - r_{t+j}^{(n-j)})]}^{\text{News about future term premia}}. \quad (8)$$

In other words, unexpected changes in long-dated forward rates must reflect either news about expected short rates in the distant future or news about future term premia. This is similar to the Campbell (1991) observation that unexpected stock returns must either be due to cash flow news or discount-rate news.

3.1. Forecasting regressions

To develop a test of whether movements in distant forward rates reflect news about future short rates or news about future term premia, we run regressions in which we use three-month changes in the forward rate, $f_t^{(n)} - f_{t-1/4}^{(n)}$, to forecast subsequent changes in forward rates over a 12-month horizon, $f_{t+1}^{(n-1)} - f_t^{(n)}$. Because $f_{t+1}^{(n-1)} - f_t^{(n)} = -(r_{t+1}^{(n)} - r_{t+1}^{(n-1)})$, this is directly equivalent to a test of Eq. (8). That is, if movements in forward rates were informative only about future short rates and not about excess bond returns, there would be no predictable mean reversion in forward rates. Conversely, if we do find evidence of mean reversion in forward rates, this maps into a particular trading strategy that earns excess returns. For example, if the ten-year forward rate jumps today and is expected to fall back over the next year, this is the same as saying that ten-year bonds are expected to outperform nine-year bonds over the next year.

We face an important data limitation in this forecasting exercise. Ideally, we would like to do everything in real terms, because our focus thus far has been on real rates. However, given the short span of the TIPS data and the fact that we are working with 12-month returns, this leaves only a dozen fully independent observations. Therefore, we worry about relying solely on TIPS forecasting regressions from 1999 to 2011.

We first focus on the nominal data, which allow us to consider a longer sample. We restrict attention to the post-1987 (post-Paul Volcker) period in which inflation expectations have been relatively well anchored in the US. It seems plausible to use the nominal data as a proxy for the missing real data over this period. In particular, our key independent variable is the change in the ten-year forward rate on FOMC announcement days. For the post-1999 period for which we have data on both, the correlation between the real and the nominal versions of this variable is 0.77. And the correlation between the nominal and real versions of our dependent variable, the change in forwards over 12-month intervals, is 0.82. This suggests that using nominal data in place of real data to extend the sample is a reasonable way to proceed.

Panel A of Table 6 presents these forecasting results. In Column 1, we begin by estimating the univariate regression

$$f_{t+1}^{\$(9)} - f_t^{\$(10)} = a + b(f_t^{\$(10)} - f_{t-1/4}^{\$(10)}) + \varepsilon_{t+1}^{\$(10)}. \quad (9)$$

That is, we use the change in the ten-year nominal forward rate over the prior quarter to predict the change in forward rates over the following 12 months.[15] Again, as a benchmark, one would expect $b=0$ under the expectations hypothesis. (Because $f_t^{\$(10)} = E_t[f_{t+1}^{\$(9)}]$ under the expectations hypothesis, it should be impossible to forecast $f_{t+1}^{\$(9)} - f_t^{\$(10)} = f_{t+1}^{\$(9)} - E_t[f_{t+1}^{\$(9)}]$.) We obtain $b=-0.343$ ($t=-3.21$), implying that a 100 bp rise in the nominal forward rate in a given quarter is associated with a 34 bp decline over the following 12 months.

In Column 2 we present instrumental variables (IV) estimates of Eq. (9) using $y_t^{\$(2)} - y_{t-1/4}^{\$(2)}$ as an instrument for $f_t^{\$(10)} - f_{t-1/4}^{\$(10)}$. These IV estimates enable us to examine the reversion following movements in forward rates that are themselves a response to changes in short rates. The large IV estimates suggest that the response of forwards to changes in short rates is quickly reverted away. The IV estimate of $b=-1.078$ ($t=-2.13$) implies that the initial response is completely reversed within 12 months. Thus, the IV estimates are consistent with the idea that the response of distant forwards to short rates primarily reflects movements in term premia as opposed to changes in expected short rates.

Columns 3 and 4 show that similar results hold when we control for the forward rate spread $f_t^{\$(10)} - y_t^{\$(1)}$ (i.e., the difference between the ten-year forward rate and the short rate) as in Fama and Bliss (1987). Our results also hold up if we control for other bond forecasting variables, including the term spread as in Campbell and Shiller (1991) or linear combinations of forward rates as in Cochrane and Piazzesi (2005) and Cieslak and Povala (2013).

In Column 5 we break down the change in the ten-year forward rate into the component that occurs on FOMC days and the component that occurs on other non-FOMC days, and we use these separately as predictive variables:

$$f_t^{\$(10)} - f_{t+1}^{\$(9)} = a + b(f_t^{\$(10)} - f_{t-1/4}^{\$(10)})_{FOMC} + c(f_t^{\$(10)} - f_{t-1/4}^{\$(10)})_{NONFOMC} + \varepsilon_{t+1}^{\$(10)}. \quad (10)$$

This approach is more tightly connected to our earlier findings, as it allows us to focus on those changes in forward rates that are associated with monetary policy announcements. The cost is that it sacrifices considerable statistical power, given the small number of FOMC days.

As shown in Column 5, the coefficient on the FOMC days part of the forward rate change ($b=-0.564$, with a t-statistic of -1.69) is somewhat larger than its counterpart for non-FOMC days ($c=-0.321$, with a t-statistic of -2.86). The IV estimates, in which we instrument for $(f_t^{\$(10)} - f_{t-1/4}^{\$(10)})_{FOMC}$ and $(f_t^{\$(10)} - f_{t-1/4}^{\$(10)})_{NONFOMC}$ with $(y_t^{\$(2)} - y_{t-1/4}^{\$(2)})_{FOMC}$ and

[15] The regressions are estimated with monthly data, so each month we are forecasting the excess return over the following 12 months. To deal with the overlapping nature of returns, t-statistics are based on Newey and West (1987) standard errors allowing for serial correlation at up to 18 lags.

Table 6
Mean reversion in forward rates.
This table forecasts of 12-month changes in forward rates using changes in forward rates over the past three months:

$$f_{t+1}^{X(9)} - f_t^{X(10)} = a + b(f_t^{X(10)} - f_{t-1/4}^{X(10)}) + \mathbf{c}'\mathbf{x}_t + \varepsilon_{t+1}^{X(10)},$$

for $X=\$$ and *TIPS*. The regressions are estimated with monthly data. To deal with the overlapping nature of the 12-month returns t-statistics are based on Newey and West (1987) standard errors allowing for serial correlation at up to 18 lags. We estimate these regressions with and without controlling for forward rate spread, $(f_t^{X(10)} - y_t^{X(1)})$. The table shows ordinary least squares (OLS) and instrumental variables (IV) estimates. In the IV specifications, we instrument for $(f_t^{X(10)} - f_{t-1/4}^{X(10)})$ using the change in nominal short rates over the past three months, $(y_t^{\$(2)} - y_{t-1/4}^{\$(2)})$. We decompose the change in ten-year forwards into components that occurred on Federal Open Markets Committee (FOMC) days and on all other days: $(f_t^{X(10)} - f_{t-1/4}^{X(10)}) = (f_t^{X(10)} - f_{t-1/4}^{X(10)})_{FOMC} + (f_t^{X(10)} - f_{t-1/4}^{X(10)})_{NONFOMC}$. In IV versions of these regressions, we instrument for $(f_t^{X(10)} - f_{t-1/4}^{X(10)})_{FOMC}$ and $(f_t^{X(10)} - f_{t-1/4}^{X(10)})_{NONFOMC}$ with $(y_t^{\$(2)} - y_{t-1/4}^{\$(2)})_{FOMC}$ and $(y_t^{\$(2)} - y_{t-1/4}^{\$(2)})_{NONFOMC}$. Panel A shows nominal forecasting results from 1987 to February 2012. Panel B shows real forecasting results from 1999 to February 2012.

	OLS (1)	IV (2)	OLS (3)	IV (4)	OLS (5)	IV (6)	OLS (7)	IV (8)
Panel A: Forecasting changes in nominal forward rates, 1987+								
$f_t^{\$(10)} - f_{t-1/4}^{\$(10)}$	−0.343 [−3.21]	−1.078 [−2.13]	−0.284 [−2.47]	−1.228 [−2.53]				
$\left(f_t^{\$(10)} - f_{t-1/4}^{\$(10)}\right)_{FOMC}$					−0.564 [−1.69]	−1.369 [−0.79]	−0.561 [−1.64]	−1.774 [−1.17]
$\left(f_t^{\$(10)} - f_{t-1/4}^{\$(10)}\right)_{NONFOMC}$					−0.321 [−2.86]	−1.084 [−2.04]	−0.256 [−2.10]	−1.238 [−2.39]
$f_t^{\$(10)} - y_t^{\$(1)}$			−0.137 [−3.15]	−0.112 [−2.12]			−0.138 [−3.21]	−0.112 [−2.10]
Constant	−0.281 [−2.76]	−0.304 [−2.59]	0.054 [0.37]	−0.037 [−0.19]	−0.283 [−2.78]	−0.308 [−2.61]	0.055 [0.37]	−0.042 [−0.21]
Number of observations	289	289	289	289	289	289	289	289
R^2	0.04		0.13		0.04		0.14	
Panel B: Forecasting changes in real forward rates, 1999+								
$f_t^{TIPS(10)} - f_{t-1/4}^{TIPS(10)}$	−0.514 [−3.99]	−0.728 [−1.27]	−0.522 [−5.21]	−0.811 [−1.88]				
$\left(f_t^{TIPS(10)} - f_{t-1/4}^{TIPS(10)}\right)_{FOMC}$					−0.567 [−2.05]	−1.081 [−1.75]	−0.533 [−2.03]	−0.590 [−1.11]
$\left(f_t^{TIPS(10)} - f_{t-1/4}^{TIPS(10)}\right)_{NONFOMC}$					−0.498 [−3.23]	−0.759 [−1.38]	−0.519 [−4.00]	−0.792 [−1.95]
$f_t^{TIPS(10)} - y_t^{TIPS(1)}$			−0.159 [−3.75]	−0.159 [−3.75]			−0.159 [−3.68]	−0.160 [−3.70]
Constant	−0.228 [−2.85]	−0.234 [−2.70]	0.164 [1.54]	0.158 [1.44]	−0.229 [−2.88]	−0.240 [−2.70]	0.164 [1.49]	0.163 [1.39]
Number of observations	142	142	142	142	142	142	142	142
R^2	0.12		0.37		0.12		0.37	

$(y_t^{\$(2)} - y_{t-1/4}^{\$(2)})_{NONFOMC}$, respectively, also result in a larger coefficient for the FOMC days piece than the non-FOMC days piece. Although the statistical significance of the FOMC days piece is marginal, the point estimates suggest that movements in forward rates on FOMC days contain just as much and perhaps even slightly more discount rate news as those on non-FOMC days.

Panel B of Table 6 presents the real analogs to Eqs. (9) and (10) for the post-1999 period. Reassuringly, we obtain similar point estimates using the TIPS data over this shorter sample period.[16] For example, the coefficient on the change in the real forward rate on FOMC announcement days from the OLS regression in Column 5 is −0.567 (t=−2.05) as compared with a value of −0.564 in the nominal data using data back to 1987.

Whether the sample period is post-1987 or post-1999, any attempt to forecast annual bond returns with a relatively small number of independent observations should be viewed with a healthy dose of skepticism. At the same time, it is important to be clear on the competing theories that are at play in this case and how they could shape one's priors. Often, when one is trying to predict asset returns, the null of no predictability has a strong ex ante theoretical standing, and so it could make sense to set a high bar for rejecting the null. But, in this case, recall that a null of no predictability is equivalent to the proposition that monetary policy shocks have a powerful effect on expected real rates ten years into the future; in other words, monetary policy is non-neutral over very

[16] One wrinkle is that we do not observe the short-term real rate, which is needed to compute the forward rate spread for long-term real bonds. Following Pflueger and Viceira (2011), we estimate the short-term real rate as the fitted value from a regression of the realized real bill return on a number of covariates.

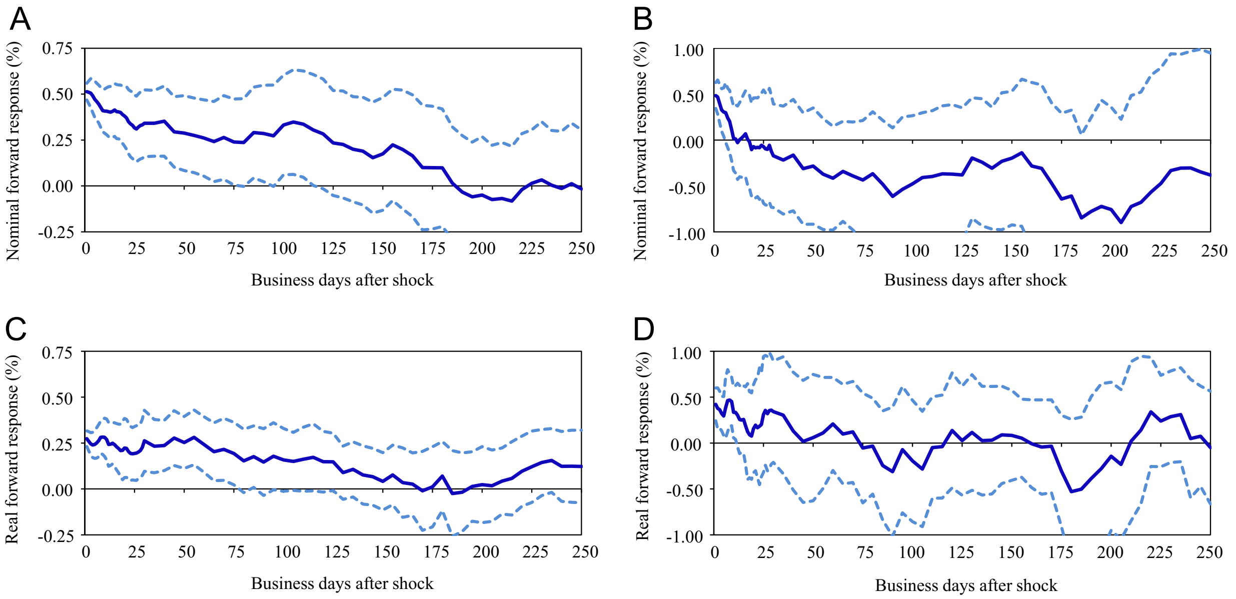

Fig. 4. Impulse response of ten-year US forwards to short-term nominal rates. Panel A plots the coefficient $b_{\$}(k)$ from estimating Eq. (11), using all days from 1987 to February 2012:

$$f_{t+k}^{\$(10)} - f_{t-1}^{\$(10)} = a_{\$}(k) + b_{\$}(k)(y_{t+1}^{\$(2)} - y_{t-1}^{\$(2)}) + \Delta\varepsilon_{t+k}^{\$(10)},$$

for horizons $k=1,\ldots,250$ days. Panel B plots repeats the same exercise, restricting attention to onlyFederal Open Markets Committee (FOMC) dates from 1987 to February 2012. Panel C plots the coefficient $b_{TIPS}(k)$ from estimating Eq. (12) on all days from 1999 to February 2012:

$$f_{t+k}^{TIPS(10)} - f_{t-1}^{TIPS(10)} = a_{TIPS}(k) + b_{TIPS}(k)(y_{t+1}^{\$(2)} - y_{t-1}^{\$(2)}) + \Delta\varepsilon_{t+k}^{TIPS(10)}.$$

Panel D plots repeats this exercise, restricting attention to only FOMC dates from 1999 to February 2012. Confidence intervals, based on Newey and West (1987) standard errors to account for the overlapping nature of the variables, are shown as dashed lines. In Panels A and C, we allow for serial correlation at up to $1.5 \times k$ lags (to the nearest integer). In Panels B and D, we allow for serial correlation between FOMC meetings at up to $1.5 \times (k/25)$ lags, because FOMC meetings occur roughly every 25 business days on average. (A) Nominal forwards, 1987+, all days. (B) Nominal forwards, 1987+, FOMC only. (C) Real forwards, 1999+, all days. (D) Real forwards, 1999+, FOMC only.

long horizons. For somebody who finds such a proposition hard to swallow, our forecasting results offer an alternative interpretation that could be more palatable, even if the statistical significance of these results is not overwhelming.

3.2. Impulse response functions

Another way to illustrate the mean reversion of forward rates is to examine their impulse response to an initial shock to short rates. To do this, we again work with daily data and the ten-year instantaneous forward rate. We begin by separately estimating

$$f_{t+k}^{\$(10)} - f_{t-1}^{\$(10)} = a_{\$}(k) + b_{\$}(k)(y_{t+1}^{\$(2)} - y_{t-1}^{\$(2)}) + \Delta\varepsilon_{t+k}^{\$(10)}, \tag{11}$$

for $k=1,\ldots,250$, using all days in the sample. That is, we regress the cumulative change in ten-year nominal forwards from day $t-1$ to day $t+k$ on the change in short-term rates from $t-1$ to $t+1$. These 250 regressions differ only in terms of the left-hand-side variable, namely, the horizon over which we compute the cumulative subsequent change in ten-year forwards.

Panel A of Fig. 4 plots the coefficient $b_{\$}(k)$ from estimating Eq. (11) on all days (i.e., FOMC and non-FOMC) using data back to 1987. The graph shows that a 100 bp shock to short-term nominal rates generates a 51 bp impulse to ten-year nominal forwards upon impact, i.e., for $k=1$. This effect is then gradually reverted away over the following nine months, consistent with the idea that the initial response reflects a change in the term premium as opposed to news about short rates. Panel B repeats the same exercise, restricting attention to only FOMC announcement dates. Consistent with our prior findings, the picture suggests that an FOMC day impulse to forward rates dissipates especially rapidly. However, as shown by the wide confidence intervals in Panel B, the standard errors increase by a factor of three or four when we focus on just FOMC days, so any inferences about the exact timing of the mean reversion are necessarily tentative in this case.

We next turn to the TIPS data for the post-1999 period and estimate

$$f_{t+k}^{TIPS(10)} - f_{t-1}^{TIPS(10)} = a_{TIPS}(k) + b_{TIPS}(k)(y_{t+1}^{\$(2)} - y_{t-1}^{\$(2)}) + \Delta\varepsilon_{t+k}^{TIPS(10)}. \tag{12}$$

Panel C plots the coefficient $b_{TIPS}(k)$ from estimating Eq. (12) on all days. The results in Panel C show that, averaging across all days, a 100 bp shock to nominal short rates is associated with a 27 bp increase in the ten-year real forward upon impact, which gradually dissipates over the following nine months. Finally, Panel D does the same thing, but focusing only on FOMC announcement dates. The initial impulse upon impact is 42 bps, which is, by construction, the same as our baseline estimate from Table 2. As in Panel B, the point estimates make it appear that this effect is largely reverted away in just a few months, but, again, the large standard errors associated with restricting attention to just FOMC days preclude precise inferences.

4. Why does monetary policy move real term premia?

We began by showing that monetary policy shocks are associated with large changes in distant real forward rates and by arguing that this association is likely to be causal in nature. Next, we showed that these changes in distant forward rates appear to reflect variation in term premia as opposed to changes in expected future short rates. This leaves a fundamental question: what is the economic mechanism by which innovations to monetary policy influence real term premia?

Broadly speaking, one can tell two types of stories. The first appeals to the standard consumption-based asset pricing model in which the real term premium is pinned down by the covariance between real bond returns and the marginal utility of the representative investor. We discuss this theory below and argue that it is unlikely to explain our results.

An alternative class of models is one in which markets are partially segmented, and term premia are determined by supply and demand effects. This is how most observers have thought about the effects of the Fed's recent quantitative easing policies, for example. These models are somewhat institutional by nature, so one can imagine many variations on the basic theme. For concreteness, we develop a particular supply and demand story based on a set of investors who care about the current yield on their portfolios. When short-term rates are low, these investors reach for yield by purchasing long-term bonds, which pushes down long-term real forward rates and lowers the term premium. We then provide some evidence that is consistent with the existence of this reaching-for-yield channel.

4.1. *Real term premia in a consumption-based asset pricing model*

According to the standard consumption-based asset pricing model, the expected excess return on long-term real bonds at time t is given by

$$E_t\left[R_{L,t+1}-R_{F,t}\right]=Corr_t\left[R_{L,t+1},-M_{t+1}\right]\sigma_t\left[R_{L,t+1}\right]\frac{\sigma_t[M_{t+1}]}{E_t[M_{t+1}]}, \tag{13}$$

where the real stochastic discount factor (SDF), M_{t+1}, depends on the marginal utility of a diversified representative investor. In light of Eq. (13), there are three ways to explain the finding that the real term premium falls when the Fed eases.

First, unexpected shifts in monetary policy could affect the volatility of bond returns $\sigma_t[R_{L,t+1}]$. However, to explain our results using this mechanism, one would further need to argue that a surprise easing lowers conditional volatility meaningfully, whereas a surprise tightening raises conditional volatility. Such an asymmetry seems difficult to motivate a priori, and little evidence exists for it in the data.[17]

Second, shifts in monetary policy could impact $Corr_t[R_{L,t+1}, -M_{t+1}]$. On the nominal side, Campbell, Sunderam, and Viceira (2013) argue that the correlation between inflation and the real SDF could vary over time, so this term could play a role in explaining time variation in inflation risk premium. It is less clear why the correlation between real bond returns and the real SDF would vary and, particularly, why it would vary meaningfully at high frequencies in response to FOMC announcements.

Finally, consider explanations that involve changes in $\sigma_t[M_{t+1}]$, the mechanism that generates time-varying risk premia in most modern consumption-based models. These models, including habit formation (Campbell and Cochrane, 1999), long-run risks (Bansal and Yaron, 2004), and time-varying disaster risk (Gabaix, 2012), share a common reduced form: $\sigma_t[M_{t+1}]$ is high during bad economic times and low during good times. However, for them to be relevant for our purposes, one would have to believe that changes in the stance of monetary policy actively cause instead of simply respond to changes in things such as long-run disaster probabilities. This seems like something of a stretch.

4.2. *A supply and demand model with yield-oriented investors*

An alternative explanation for why monetary policy can move term premia is based on supply and demand effects that operate in partially segmented bond markets. We illustrate this point with a simple model featuring a set of investors who care about the current yield on their portfolios. The key assumptions of the model are as follows. There are two dates, 1 and 2. The real log short rate at time 1, r_1, is set by the central bank. The real log short rate at time 2, r_2, is initially uncertain. Moreover, monetary policy is assumed to be neutral in the long run. Thus, both $E[r_2]$ and $Var[r_2]$ are outside of the time 1 control of the central bank and should be thought of as pinned down by long-run macroeconomic fundamentals. The only endogenous variable is y_2, the time 1 log yield on real long-term (two-period) bonds, and our interest is in seeing how y_2 varies with the stance of monetary policy as summarized by r_1.

A fraction α of investors are yield-oriented with non-standard preferences described below, and a fraction $(1-\alpha)$ are expected return-oriented with conventional mean-variance preferences. Both investor types have unit risk tolerance.

Expected return-oriented investors have zero initial wealth and construct long-short positions to maximize $E[w_R]-Var[w_R]/2$, where w_R is their future wealth. If they purchase b_R units of long-term bonds and finance this

[17] Lee (2002) estimates generalized autoregressive conditional heteroskedasticity (GARCH) models that enable him to estimate the impact of a surprise FOMC easing separately versus a surprise tightening on interest rate volatility. While a surprise tightening has a larger impact on volatility than a surprise easing, the evidence suggests that both positive and negative surprises raise volatility.

position by rolling over short-term borrowing, their future wealth is $w_R = b_R(2y_2 - r_1 - r_2)$.[18] Thus, expected return-oriented investors solve

$$\max_{b_R}\left\{b_R(2y_2 - r_1 - E[r_2]) - b_R^2\, Var[r_2]/2\right\}, \tag{14}$$

and their demand for long-term bonds is

$$b_R(y_2) = (Var[r_2])^{-1}(2y_2 - r_1 - E[r_2]). \tag{15}$$

By contrast, yield-oriented investors pick their holdings of long-term bonds b_Y, to solve

$$\max_{b_Y}\left\{b_Y \cdot (2y_2 - 2r_1) - b_Y^2\, Var[r_2]/2\right\}. \tag{16}$$

The only difference between Eqs. (16) and (14) is that in Eq. (16) we have replaced $E[r_2]$ in the first term with r_1. The interpretation is that yield-oriented investors care about the spread in current yield between long- and short-term bonds (as captured by $2y_2 - 2r_1$) as opposed to the spread in expected returns (as captured by $2y_2 - r_1 - E[r_2]$). Said differently, if the yield curve is upward-sloping simply because $E[r_2]$ exceeds r_1, long-term bonds would be more attractive to the yield-oriented investors but not to the expected return-oriented investors. Thus, the demand for long-term bonds from yield-oriented investors depends on the difference in current income from owning long- versus short-term bonds:

$$b_Y(y_2) = (Var[r_2])^{-1}(2y_2 - 2r_1). \tag{17}$$

We assume there is a fixed supply Q of long-term real bonds. The market clearing condition for long-term bonds is $Q = \alpha\, b_Y(y_2^*) + (1-\alpha)b_R(y_2^*)$, which implies that the equilibrium long-term forward rate is

$$\overbrace{2y_2^* - r_1}^{\text{Forward rate}} = \overbrace{E[r_2]}^{\text{Expected short rate}} + \overbrace{Q\, Var[r_2] - \alpha(E[r_2] - r_1)}^{\text{Term premium}}. \tag{18}$$

Similarly, the expected excess return on long-term bonds is

$$2y_2^* - E[r_2] - r_1 = \overbrace{Q\, Var[r_2]}^{\text{Traditional term premium}} - \overbrace{\alpha(E[r_2] - r_1)}^{\text{Reaching-for-yield term premium}}. \tag{19}$$

Eqs. (18) and (19) show that the term premium has two components. A traditional component, $Q\, Var[r_2]$, depends on bond supply and fundamental uncertainty, and a reaching-for-yield component, $-\alpha(E[r_2] - r_1)$, depends on the fraction of yield-oriented investors and the level of short-term interest rates.

The reaching-for-yield term in Eqs. (18) and (19) is what enables the model to rationalize our prior findings, namely, that an easing of monetary policy is associated with a decline in distant real forwards and a decline in the real term premium. When the central bank cuts the short rate, $E[r_2] - r_1$ rises and the term premium falls. Intuitively, this is because yield-oriented investors are hungrier for current income when r_1 is low. As a result, they are willing to take on more duration risk by purchasing higher-yielding long-term bonds. And due to the limited risk tolerance of investors on the other side of the trade, this shift in demand lowers the term premium on these long-term bonds. This explanation draws no distinction between movements in rates on FOMC versus non-FOMC days. It does not matter whether rates move due to news about the Fed's reaction function or news about macroeconomic fundamentals. Yield-oriented investors care about the differential current yield from holding long-term bonds irrespective of its root cause.

Why, according to this view, would one expect this lower term premia to accrue largely over the following 12 months? There are a few possibilities. A decline in short rates could only temporarily boost demand for long-term Treasuries from yield-oriented investors. Perhaps some yield-oriented investors initially respond to a drop in short rates by taking on more duration risk, but over time they instead shift toward taking on more credit risk. Alternatively, if arbitrage capital moves slowly in response to changes in risk-adjusted returns, the demand shock from yield-oriented investors could be met with increased arbitrageur capital over time. Or, following Greenwood, Hanson, and Stein (2010), the increased demand for long-term bonds could be gradually accommodated by non-financial firms that adjust their debt maturity in response to shifts in investor demand.

4.3. Evidence on the behavior of yield-oriented investors

In addition to rationalizing the movements in real forward rates and term premia shown in Sections 2 and 3, the model offers an additional set of predictions. Specifically, if we can identify a priori those investors who are most prone to be yield oriented, their holdings of long-term bonds should be increasing in the yield spread. This follows immediately from Eq. (17), which says that the demand of yield-oriented investors is a function of $(y_2 - r_1)$.

The holdings of the investing public as a whole must equal the fixed supply of long-term bonds, so there must be other investors (e.g., broker-dealers or hedge funds) who care less about current yield differentials and more about expected returns and who take the other side of the trade. In what follows, we use commercial banks as a proxy for yield-oriented investors and primary dealers as a proxy for expected return-oriented investors. The logic of this split is based on existing accounting conventions, which should arguably have the effect of making banks more concerned with current yield than dealers.

4.3.1. Commercial banks

We use quarterly Call Report data on the duration of commercial bank security portfolios to test the hypothesis that banks act like the yield-oriented investors in our model. Several factors suggest that commercial banks could be prone to behave in a yield-oriented fashion. First, the vast majority of commercial banks (weighted by market value) are publicly traded, so bank managers with short horizons could be tempted to take actions that boost current reported earnings at the expense of longer-term

[18] We work with log returns to facilitate the exposition. This can be viewed as a linear approximation to a model based on simple returns. Specifically, the two-period excess simple return on a strategy of buying two-period bonds, financed by rolling over short-term debt, is $(1+Y_2)^2 - (1+R_1)(1+R_2) \approx 2y_2 - r_1 - r_2$, where $x = \ln(X)$.

earnings (Stein, 1989).[19] Second, due to GAAP accounting conventions, a bank can typically boost near-term accounting earnings simply by replacing low-yielding securities in its non-trading accounts with higher-yielding securities.[20] This is because interest income on non-trading account securities flows through the income statement, but unrealized gains and losses on such securities do not flow through income. Thus, a desire to boost current reported profits could lead bank managers to invest more aggressively in long-term securities when the yield curve is steep. Finally, because GAAP earnings also drive changes in regulatory capital, a bank could boost its capital ratios and generate regulatory slack in the near term by engaging in a larger carry trade when the curve is steep.

Given the coarse disclosure available in the Call Reports, we focus on a crude measure of securities portfolio duration: the aggregate fraction of non-trading account securities with a current remaining maturity (for fixed-rate securities) or next re-pricing date (for floating-rate securities) of one year or longer: (SEC_{LT}/SEC). This measure is available beginning in 1988. Using quarterly data, we estimate specifications of the form

$$\Delta(SEC_{LT}/SEC)_t = a + b\,\Delta(y_t^{\$(10)} - y_t^{\$(1)}) + \Delta u_t, \tag{20}$$

where ($y^{\$(10)}$–$y^{\$(1)}$) is the yield spread, measured as the difference in current yield between ten- and one-year nominal Treasuries. A finding that $b>0$ would suggest that banks reach for yield, buying more long-term bonds when the yield curve steepens.

Table 7 presents the results from this exercise. Column 1 shows a strong positive relation between $\Delta(SEC_{LT}/SEC)$ and $\Delta(y^{\$(10)}$–$y^{\$(1)})$.[21] In terms of dollar magnitudes, the coefficient in Column 1 suggests that a 100 bp decline in the short rate, holding fixed the long rate, leads to a 1.06 percentage point increase in the share of bank securities that are long term. Bank securities have averaged roughly 18.5% of total bank assets since 1988. As of 2010:Q4, commercial bank assets were \$11,728 billion, so this means that a 100 bp increase in the yield spread raises bank demand for long-term securities by \$23 billion ($= 1.06\% \times 18.5\% \times \$11{,}728$). So a 300 bp swing in the yield spread, roughly the range over a full easing cycle, would boost demand by \$69 billion. The less-than-one-year versus longer-than-one-year margin is potentially only part of the overall portfolio adjustment process. Banks could also be extending their duration within the longer-than-one-year bucket. And banks are just one set of investors who could care about current income. Thus, the results in Table 7 suggest that the induced shift in total demand from all yield-oriented investors could be substantial.

The remaining columns of Table 7 test another implication of the reaching-for-yield story. Specifically, if reaching for yield is partially driven by a desire to manage reported earnings, then this tendency should be more pronounced for publicly traded banks than for privately held banks. If reaching for yield is driven solely by a desire to boost capital ratios and maintain regulatory slack, then one would not expect to see much of a difference between public and private banks. To investigate this issue, we construct two versions of SEC_{LT}/SEC, one for public banks and another for private banks. As shown in Column 2, the results for public banks are similar to those for all banks. This is true almost by construction because a large majority of aggregate commercial banking assets are held by public banks. Consistent with the earnings-management hypothesis, Column 3 shows that yield-chasing behavior is less pronounced amongst private banks. The coefficient of 0.674 on $\Delta(y^{\$(10)}$–$y^{\$(1)})$ for private banks in Column 3 is only half of the corresponding coefficient for public banks in Column 2. However, as shown in Column 4, we cannot reject the hypothesis that the coefficient for public banks is the same as that for private banks. Thus, the split between public and private banks goes in the direction predicted by the earnings-management story, but the evidence on this front is statistically weak. Moreover, the positive albeit insignificant coefficient for private banks suggests that a desire to maintain regulatory slack could also play some role.

Another question has to do with the persistence of the shifts in banks' demands for long-term bonds. We find some tentative evidence (not reported) suggesting that these demand shocks are gradually reversed over roughly the following eight quarters. One way to see this is to add lagged values of $\Delta(y^{\$(10)}$–$y^{\$(1)})$ to Eq. (20) and then examine the cumulative sum of coefficients on contemporaneous and lagged changes in the yield spread.

4.3.2. Primary dealers

Next we examine the Treasury holdings of primary dealers. We think of primary dealers as a natural proxy for the expected return-oriented investors in our model, the arbitrageurs who accommodate demand shocks coming from yield-oriented investors. Importantly, primary dealer activities are housed either within broker-dealers or in commercial bank trading departments. As a result, unlike banks' non-trading accounts, primary dealers operate entirely on mark-to-market accounting. So, even if they

[19] Specifically, one could extend the Stein (1989) analysis to consider an earnings management technology in which the trade-off between current earnings and expected future earnings (or future earnings risk) varies over time. In the present case, the current earnings benefit of taking interest rate risk is greater when the yield curve is steep, so one would expect managers with short horizons to take more interest rate risk at those times.

[20] Almost all non-trading account securities are treated as available-for-sale under US Generally Accepted Accounting Principles (GAAP). Although these securities are marked to market, unrealized gains and losses do not flow through the income statement (as do changes in the value of trading account securities). Instead, mark-to-market gains and losses flow through other comprehensive income and are accumulated on the balance sheet until realized. By contrast, interest income on these securities flows directly through the income statement. Thus, a bank can boost its near-term reported earnings by substituting high-yielding for low-yielding securities. A similar effect obtains for the small fraction of non-trading securities that are carried on the balance sheet on a historical cost basis (a.k.a., held-to-maturity securities). The workings of international accounting rules under International Financial Reporting Standards are broadly similar to US GAAP in this respect.

[21] We have also tried regressing $\Delta(SEC_{LT}/SEC)$ on both $\Delta y^{\$(10)}$ and $\Delta y^{\$(1)}$ separately. The coefficient on $\Delta y^{\$(10)}$ is positive and significant, and the coefficient on $\Delta y^{\$(1)}$ is negative and significant. The absolute magnitudes of the two coefficients are similar, consistent with the logic of Eq. (20).

Table 7

Duration of commercial bank securities portfolios and the yield spread.

Regressions of quarterly changes in the aggregate duration of bank securities portfolios on quarterly changes in the yield spread from 1988 to 2010:

$$\Delta(SEC_{LT}/SEC)_t = a + b\Delta(y_t^{\$(10)} - y_t^{\$(1)}) + \Delta u_t.$$

t-Statistics, based on robust standard errors, are shown in brackets. SEC_{LT}/SEC is the fraction of non-trading account securities with a remaining maturity (for fixed rate securities) or next re-pricing date (for floating rate securities) of one year or more. Column 1 shows the result for all banks. Columns 2 and 3 show results for publicly traded banks and for private banks, respectively. Finally, Column 4 shows the difference between public and private banks. Thus, the t-statistics in Column 4 enables one to test the hypothesis that the coefficients for public and private banks are equal. We classify a commercial bank as publicly traded if its parent bank holding company has a valid Center for Research in Security Prices link in the table maintained by researchers at the Federal Reserve Bank of New York. This linking table is available online at http://www.newyorkfed.org/research/banking_research/datasets.html.

	All banks (1)	Public banks (2)	Private banks (3)	Difference: Public–Private (4)
$\Delta(y^{(\$(10)}–y^{(\$(1)})$	1.060	1.229	0.674	0.555
	[3.76]	[3.07]	[1.33]	[0.75]
Constant	0.009	0.003	−0.009	0.012
	[0.08]	[0.02]	[−0.07]	[0.07]
Number of observations	91	91	91	91
R^2	0.13	0.11	0.04	0.01

Table 8

Duration of primary dealer Treasury holdings and the yield spread.

Regressions of weekly changes in the aggregate duration primary dealer Treasury holdings on weekly changes in the yield spread from July 2001 to February 2012:

$$\Delta NETDUR_t = a + b\Delta(y_t^{\$(10)} - y_t^{\$(1)}) + \Delta u_t.$$

$NETDUR_t$ is defined in Eq. (21). The even-numbered columns include controls for the weekly change in the scale of dealers' net position in Treasuries and all other reportable fixed income asset classes as well as a full set of week-of-year dummies. t-Statistics, based on robust standard errors, are shown in brackets. Columns 1 and 2 show this exercise for our baseline measure of the duration of dealers' Treasury holdings. Columns 3 and 4 repeat this exercise using a more comprehensive duration measure based on dealers' holdings of both Treasuries and agency debentures. The primary dealer data are available on-line at http://www.newyorkfed.org/markets/gsds/search.cfm.

	Treasury holdings		Treasury plus agency holdings	
	(1)	(2)	(3)	(4)
$\Delta(y^{(\$(10)}–y^{(\$(1)})$	−0.973	−0.884	−0.451	−0.401
	[−3.12]	[−3.00]	[−3.85]	[−3.53]
Constant	0.004	−0.355	0.002	−0.118
	[0.12]	[−0.94]	[0.17]	[−0.84]
Controls	No	Yes	No	Yes
Number of observations	555	555	555	555
R^2	0.03	0.15	0.04	0.20

wanted to manage their earnings, playing the term spread would be less of a sure thing for them.

We work with data from the Federal Reserve Bank of New York on the aggregate holdings of primary dealers. A key advantage of this data is that we have high frequency observations of dealer holdings by maturity. Specifically, for each week beginning in July 2001, we have data on the aggregate net (long minus short) dealer holdings of Treasury bills (all of which mature in less than one year) and nominal coupon-bearing Treasuries broken into four buckets by remaining maturity: shorter than three years, three to six years, six to 11 years, and longer than 11 years.[22]

Measuring the net duration of primary dealer positions is a bit tricky because dealers can be net short Treasuries, both in a given maturity bucket and overall. To deal with this complication, we compute

$$NETDUR_t = \frac{0.5Q_{(TB)t} + 1.5Q_{(m \le 3)t} + 4.5Q_{(3<m\le 6)t} + 8.5Q_{(6<m\le 11)t} + 14Q_{(11<m)t}}{|Q_{(TB)t}| + |Q_{(m\le 3)t}| + |Q_{(3<m\le 6)t}| + |Q_{(6<m\le 11)t}| + |Q_{(11<m)t}|}. \quad (21)$$

The numerator is proxy for the net dollar duration of dealers' Treasury holdings. Then, to scale this variable, we divide it by the sum of the absolute positions in each maturity bucket. Thus, the scaled variable is like $(DUR_t^{LONG}\,LONG_t - DUR_t^{SHORT}\,SHORT_t)/(LONG_t + SHORT_t)$.[23]

A variety of factors besides those in our model could impact the duration of dealers' Treasury holdings. In an attempt to control for some of these, we include proxies for the weekly change in the scale of dealers' net positions in Treasuries as well as the change in their net positions across all other reportable fixed income asset classes. A related concern is that high-frequency variation in the maturity structure of dealer positions could be driven by seasonal patterns of Treasury issuance; e.g., due to seasonal fluctuations in T-bill supply or the large offerings of longer-term notes and bonds in February, May, August, and November. To deal with this, we include a full set of week-of-year dummies to soak up any seasonal fluctuations in Treasury supply.

Using weekly data, we then estimate specifications of the form

$$\Delta NETDUR_t = a + b\,\Delta(y_t^{\$(10)} - y_t^{\$(1)}) + \mathbf{c}'\mathbf{x}_t + \Delta u_t. \quad (22)$$

If dealers function as the expected return-oriented investors in our model, we would expect to find $b<0$. Table 8 presents the results from this exercise. Column 1 shows a strong negative relation between $\Delta NETDUR_t$ and $\Delta(y^{\$(10)}–y^{\$(1)})$. Column 2 adds the various controls, including the

[22] While primary dealers also report their holdings of TIPS, these are not broken out by maturity and so we do not use them to compute our measure of the maturity structure of dealers' holdings. However, dealers' holdings of TIPS are small relative to their overall Treasury holdings, so this choice has little impact on the resulting measure.

[23] Alternate approaches are to work directly with the numerator of $NETDUR_t$ or to scale the numerator by the amount of outstanding Treasuries. These approaches are less desirable because the resulting measures are impacted by variation in the scale of brokers' holdings relative to market as a whole and, thus, are no longer pure maturity measures. Nevertheless, we have experimented with these constructions and generally obtain similar results.

week-of-year dummies. While the controls substantially increase the overall explanatory power of the regression, they have little impact on the coefficient of interest. Columns 3 and 4 repeat these exercises using a duration measure based on dealers' holdings of both Treasuries and agency debentures, because the latter are seen as a close substitute for Treasuries by many investors. This yields broadly similar conclusions.

How do the dollar magnitudes implied by Table 8 for dealers compare with those for commercial banks? Our estimates suggest that a 100 bp increase in the term spread raises commercial banks' demand for long-term securities by \$23 billion. For the sake of argument, assume this means that banks buy \$23 billion of ten-year Treasury notes and sell a corresponding amount of short-term T-bills. Also assume the ten-year notes have a duration of eight years and bills have zero duration. To make the comparison, we need to convert our estimates in Table 8 for broker dealers into dollar magnitudes. To do so, we note that dealers' average absolute position in Treasury securities is on the order of \$100 billion over our sample. Thus, our estimates in Column 1 of Table 8 suggest that a 100 bp increase in the term spread in a given week leads broker-dealers to sell \$12.2 billion ($=0.973 \times 100 \div 8$) ten-year notes and purchase a corresponding amount of short-term bills.

This rough calculation suggests that broker-dealers are acting as economically meaningful arbitrageurs, accommodating roughly half of the demand by yield-oriented commercial banks in the wake of a shock to the term spread. Other players besides banks also could be reaching for yield, and other investors besides broker-dealers also could be acting as arbitrageurs. So we cannot claim to have an overall handle on the magnitude of either the aggregate demand shock or the arbitrage response.

Finally, we examine the dynamics of $NETDUR_t$ following a shock to the yield spread. Specifically, we separately estimate

$$NETDUR_{t+k} - NETDUR_{t-1} = a(k) + b(k)[(y_t^{\$(10)} - y_t^{\$(1)}) - (y_{t-1}^{\$(10)} - y_{t-1}^{\$(1)})] + \Delta\varepsilon_{t+k}, \quad (23)$$

for $k=0, 1,\ldots,52$. Thus, as above, these regressions differ solely in terms of the differencing horizon on the left-hand side (the estimate for $k=0$ corresponds to the estimates in Column 1 of Table 8). Fig. 5 plots the coefficients, $b(k)$, versus horizon k. The point estimates suggest that the initial impulse to the duration of dealers' Treasury holdings persists for roughly five months but then largely vanishes within nine months. Interestingly, this roughly matches the horizon over which the impulse from short-term nominal rates to distant real forwards is reverted away in Fig. 5. One possible interpretation, in the spirit of Grossman and Miller (1988), is that primary dealers function as front-line arbitrageurs in response to a demand shock, but, over time, more arbitrage capital enters the market, allowing the dealers to unwind their positions and reversing the initial price impact. The wide confidence intervals in Fig. 5 underscore that our estimates of the timing of dealers' unwind are imprecise, so this interpretation is necessarily somewhat speculative.

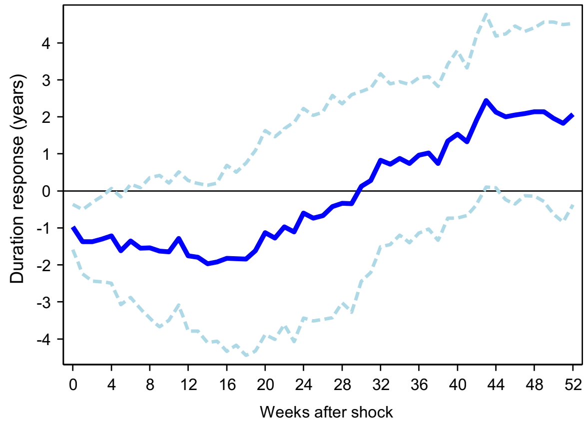

Fig. 5. Impulse response of primary dealer *NETDUR* to the yield spread. The figure plots the coefficient $b(k)$ from estimating the following regressions using weekly data from July 2001 to February 2012:

$$NETDUR_{t+k} - NETDUR_{t-1} = a(k) + b(k)[(y_t^{\$(10)} - y_t^{\$(1)}) - (y_{t-1}^{\$(10)} - y_{t-1}^{\$(1)})] + \Delta\varepsilon_{t+k},$$

for horizons $k=0, 1,\ldots,52$ weeks. Confidence intervals, based on Newey and West (1987) standard errors to account for the overlapping nature of the variables, are shown as dashed lines. We allow for serial correlation at up to $1.5 \times k$ lags (to the nearest integer).

4.4. Other supply and demand channels

Our theory of yield-oriented investors is one specific example of a supply and demand channel that connects monetary policy shocks to real term premia. However, one can tell other stories in a similar spirit. For instance, Hanson (2014) and Malkhozov, Mueller, Vedolin, and Venter (2014) argue that shifts in expected mortgage refinancing generate shocks to the aggregate supply of duration, which impact term premia. A positive shock to interest rates lowers expected mortgage refinancing, causing the duration of existing mortgage-backed securities (MBS) to extend. As a result, the quantity of interest rate risk that bond investors must bear increases following a shock to interest rates, leading the term premium to rise. Consistent with this, Hanson (2014) finds evidence that measures of aggregate MBS market duration positively forecast bond returns and that shifts in MBS duration help explain the sensitivity of distant real forwards to short-term nominal rates.

Another possible demand-side explanation is that investors have a mistaken tendency to extrapolate current short-term real rates into the distant future. Our model, in which some investors are yield-oriented due to agency effects, is isomorphic to one in which some investors have highly extrapolative beliefs, assuming that the future short rate will be the same as today's. Consistent with this, Cieslak and Povala (2014) argue that investors make systematic expectational errors about the near-term path of real short rates. Because we focus on distant forward real rates, a behavioral explanation for our results would need to invoke more severe mistakes. Investors would need to think that current policy has a large impact on real short rates more than ten years into the future. However, such a belief could be consistent with the theory of natural expectations developed by Fuster, Laibson, and Mendel (2010), in which expectations are a combination of highly extrapolative intuitive expectations and traditional rational expectations. And, consistent with this, Piazzesi, Salomao, and Schneider (2013) find that forecasters over-estimate the persistence of both the level and the slope of the yield curve.

5. Conclusions

Changes in the stance of monetary policy have a surprisingly strong impact on distant forward real interest rates. These movements in forward rates appear to reflect changes in term premia, which largely accrue over the next year, as opposed to varying expectations about future real rates. Moreover, our evidence suggests that one driving force behind time-varying term premia is the behavior of yield-oriented investors, who react to a cut in short rates by increasing their demand for longer-term bonds, thereby putting downward pressure on long-term rates.

Our work raises, but does not answer, a series of questions about the ultimate economic importance of this monetary transmission channel. In particular, suppose that a monetary easing lowers long-term real rates through the mechanism we have described. What could the resulting impact on corporate investment be? On the one hand, the fact that the effect of monetary policy on long-term real rates is transitory (i.e., it is reversed after about a year) could seem to imply that it would matter less for corporate capital-budgeting decisions. On the other hand, some firms could view the temporarily lower long-term rates as a market-timing opportunity, i.e., a window during which it is particularly attractive to issue long-term debt. This in turn could serve to stimulate their investment.[24]

While we have focused narrowly on term premia in the Treasury market, the idea that monetary policy can influence bond market risk premia has potentially broader implications. Much recent work has been motivated by the hypothesis that accommodative monetary policy can reduce credit-risk premia.[25] A promising avenue for future work would be to study these two channels of monetary transmission in a unified setting. For example, in the context of our model, one could allow yield-oriented investors to choose among not only short-term and long-term Treasuries, but also defaultable credit instruments (corporate loans, mortgages, etc.). This would presumably yield a set of predictions about the comovement of term premia and credit-risk premia in response to changes in monetary policy and could be the basis for a wider-ranging and more integrated empirical investigation of these phenomena.

[24] Similar issues arise when a firm makes investment and financing decisions in the face of a transitory equity mispricing. Stein (1996) and Baker, Stein and Wurgler (2003) argue that the investment of financially constrained firms is most sensitive to equity mispricing. One could conjecture that the same would be true here, that is, constrained firms would be most likely to adjust investment in response to policy-induced changes in term premia. At the same time, the financing (i.e., debt maturity) decisions of unconstrained firms would adjust more as in Greenwood, Hanson, and Stein (2010). However, unlike constrained firms, this would not have a large impact on their investment.

[25] See, e.g., Rajan (2005), Borio and Zhu (2008), Greenwood and Hanson (2013), Jiménez, Ongena, Peydró, and Saurina (2014), Adrian and Shin (2010), Landier, Sraer, and Thesmar (2011), and Gertler and Karadi (2013).

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