Risk Premia in Gold Lease Rates*

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Preliminary, Comments Welcome

Abstract

Gold is an important global reserve asset, widely held by central banks. In this paper, we study a measure of the opportunity costs of holding gold: the *gold lease rates* – interests paid in gold for borrowing gold. We find that gold lease rates are economically significant in magnitude. Expected excess returns of "gold bonds" are mostly positive, suggesting that they are perceived as risky investments. These risk premia are highly time-varying and strongly increasing in the slope of the gold yield curve, leading to a decisive rejection of the expectations hypothesis. In addition to the slope, gold volatility and factors in equity, credit, and currency markets also predict excess returns of gold bonds.

Keywords: gold lease rates, expectation hypothesis, risk premia, term structure
JEL codes: G12

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1 Introduction

Gold is an important global reserve asset. According to World Gold Council (2013a), as of the second quarter of 2012, gold accounts for about 13%, equivalent to $1.56 trillion, of the $12 trillion total reserves held by central banks. This amount makes gold the third largest reserve asset globally, only behind the U.S. dollar and the Euro, but larger than the British Pound, Japanese Yen, and Swiss Franc. In particular, the United States (the Euro area) holds 69.8% (56.0%) of their foreign exchange reserves in gold.[1] In three joint statements issued in 1999, 2004, and 2009, fifteen central banks led by the ECB have affirmed that "gold will remain an important element of global monetary reserves."[2]

Gold also plays an increasing role as financial collateral, formally recognized so by the Basel Accords, an international standard for regulating bank capital (see Basel II 2006 and Basel III 2012). Leading exchanges and clearinghouses, such as CME Group, ICE Clear Europe, and LCH.Clearnet, accept gold as eligible collateral for margining purposes.[3] J.P. Morgan, a major tri-party repo collateral manager, also allows clients to post gold collateral to satisfy security lending and repo obligations.[4]

The status of gold as a reserve asset and collateral raises important questions: What are the opportunity costs of holding gold or pledging gold as collateral, over short and long terms? Equivalently, what is the term structure of interest rates for borrowing and lending gold? And what determines the risk premium associated with gold borrowing and lending?

In this paper, we study the dynamics and risk premia of "gold interest rates" – *the interest rates paid in gold for borrowing gold* – often known as the gold lease rates. For example, an investor lending 100 ounces of gold at a lease rate of 2% per annum will earn 102 ounces of gold after a year. The gold lease rates are, therefore, analogous to interest rates of the U.S. dollar. Just as the latter measure the costs of dollar borrowing, the former capture the returns – in terms of gold – earned by central banks and investors for lending gold held as part of their reserves and asset allocation. Equivalently, the gold lease rates capture the opportunity costs for central banks and investors for holding gold passively without lending

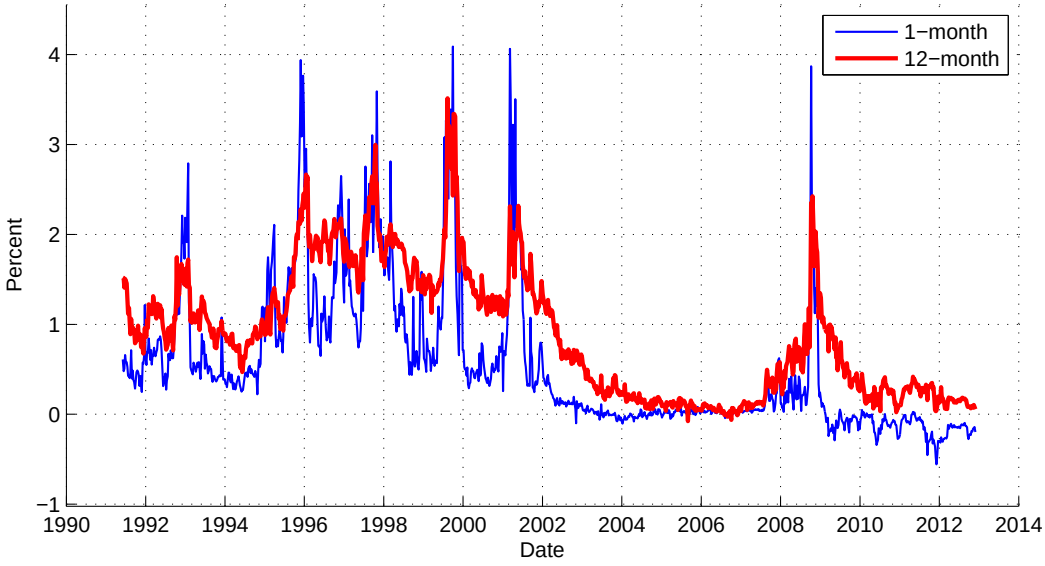
[1] See World Gold Council (2013b). For the Euro area, the largest holders include Germany (66.3%), Italy (64.9%), France (64.5%), Netherlands (52%), and Portugal (84.2%).

[2] See http://www.ecb.eu/press/pr/date/1999/html/pr990926.en.html , http://www.ecb.eu/press/pr/date/2004/html/pr040308.en.html, and http://www.ecb.eu/press/pr/date/2009/html/pr090807.en.html.

[3] See http://www.cmegroup.com/trading/metals/cme-group-accepts-gold-as-collateral.html for the CME announcement, http://ir.theice.com/releasedetail.cfm?ReleaseID=527772 for the ICE announcement, and http://www.lchclearnet.com/member_notices/circulars/2012-08-21.asp for the LCH.Clearnet announcement.

[4] See "J.P. Morgan Will Accept Gold as Type of Collateral," *Wall Street Journal*, February 8, 2011.

Figure 1: Time series of gold lease rates from June 1991 to December 2012.



it, or for pledging gold as collateral.

Embedded in the gold lease rates are risk premia required by investors for taking gold lending risks. As is evident in Figure 1, gold lease rates display substantial variations over time.[5] Characterizing the risk premia in gold lease rates, as well as comparing them to the risk premia in U.S. dollar interest rates at comparable maturities, are the primary objective of our study.

Our theoretical benchmark for studying risk premia in gold lease rates is the expectation hypothesis (EH). The EH is one of the most established theories of interest rates. Applied to the gold lending market, the EH posits that excess returns on gold bonds are not predictable, or equivalently, the risk premium in the gold lease rates is constant. Moreover, although the EH has been rejected by many studies using long-term interest rates of major fiat currencies,[6] evidence regarding the EH for *short-term* interest rates is far from conclusive. For example, Longstaff (2000) and Downing and Oliner (2007) find strong support for the EH in the U.S. repo market and U.S. commercial paper market, whereas Buraschi and Menini (2002) and Piazzesi and Swanson (2008) document violations of the EH for the German repo "specialness"

[5]During the recent financial crisis, short term gold lease rates occasionally become negative. In order to mitigate effects of possible abnormal market behaviors during these episodes, in our main analysis we start the data sample in June 1991 and end it in August 2007, right before the financial crisis.

[6] For U.S. evidences, see Dai and Singleton (2002) and references therein. For international evidences, see Jotikasthira, Le, and Lundblad (2013) and references therein.

and U.S. Fed Fund futures market. Given these mixed findings on short-term interest rates of the U.S. dollar and the Euro, a natural question is whether the EH accurately describes the joint dynamics of gold lease rates at different maturities.

Using available gold lease rates up to one year in maturity,[7] sampled weekly from June 1991 to August 2007, we find that the EH is decisively rejected in the gold leasing market. That is, risk premia in the term structure of gold lease rates are strongly time-varying, even at short maturities. By sharp contrast, the EH cannot be consistently rejected for Eurodollar deposit rates at the same maturities.[8]

We characterize the time-varying risk premia in the gold leasing market through a series of tests. Our first step is the classical Campbell and Shiller (1991) regression. For an investor who lends gold for a longer term (e.g. 12-month) by borrowing gold for a shorter term (e.g. 1-month), the Campbell and Shiller regression tests whether the positive yield spread earned by the investor moves one-for-one with the mark-to-market loss on the long-term loan due to yield changes. If the EH holds, regressing the mark-to-market loss on the yield spread should give a coefficient of one. In the gold leasing markets, we find that the regression coefficients are all negative, ranging from about 0 for the 8-week maturity to -1.5 for the 52-week maturity. All these coefficients are significantly below one, which constitutes strong evidence against the EH. The same regression applied to Eurodollar rates of the same maturities, however, gives coefficients that are all statistically indistinguishable from one. That is, the EH cannot be rejected in the Eurodollar data.

Our second test is the Cochrane and Piazzesi (2005) regression, which uses all yields, not only the long-short spread, to predict excess returns in bond markets. We find that the gold lease rates strongly predict monthly excess returns on gold bonds, with adjusted R^2's between 12.1% and 25.4%. By contrast, the same regressions applied to the Eurodollar data only yield adjusted R^2's between 3.4% and 6.5%, with borderline or insignificant coefficients.

Finally, we estimate a two-factor no-arbitrage dynamic term structure model of the gold lease rates, applying the framework of Joslin, Singleton, and Zhu (2011) (JSZ). Relative to regressions, a no-arbitrage model better links excess returns across different maturities and, in combination with Kalman filtering, allows a more robust treatment of potential

[7]Since no central bank sets the target interest rate on gold bonds, gold lease rates are determined in the marketplace and are published by the London Bullion Market Association for 1-, 2-, 3-, 6-, and 12-month maturities.

[8]The Eurodollar deposit rates are for deposits in U.S. dollars outside of the U.S. These rates serve as a natural comparison with the gold lease rates because both rates are set outside the U.S. and have similar credit risk and overlapping maturities. Note that the Eurodollar deposit rates are not the same as the Eurodollar futures contracts traded on the CME.

measurement errors. We find that the estimated parameters governing the market price of risk are statistically and economically significant for the gold lease rates, implying a time-varying risk premium and, again, a resounding rejection of the EH. The same term structure model applied to the Eurodollar data produces insignificant market price of risk parameters and fails to reject the EH. In particular, risks about both the level and slope of the gold lease rate curve are significantly priced, and the risk premia are strongly increasing in the current slope. Over our sample period, the slope of the gold lease rate curve is highly time-varying and mostly positive. This suggests that lending gold is usually viewed by the market as a risky investment.

We conduct a number of robustness checks, including alternative data source (CME futures data, see Appendix A). Our main results remain unchanged; the EH is still rejected in the gold lease rates.

The strong rejection of the EH in the gold lease rates, as well as the failure to reject it for the Eurodollar data, reveals substantially different behaviors of risk premia in the two markets. A natural question is whether our results can survive the inclusion of additional factors that explain asset returns in other markets, and whether these additional variables help predict excess returns of gold bonds.

We explore these questions by incorporating a variety of factors – including macroeconomic forecasts, factors for the markets of physical gold, equities, credit, and currencies, and measures of volatility and liquidity – in a predictive regression. To accommodate the macro variables, we run this regression using monthly data. We find that after including these additional variables, the slope of gold yield curve remains a significant predictor of excess returns and even becomes economically larger. Excess returns in gold bonds are also significantly predicted by gold volatility, small-firm premium, credit spread, foreign currency returns, and a measure of the global risk of the US dollar. These results reveal that the risk premia embedded in gold lease rates also (partly) reflect those in many other segments of financial markets. In sharp contrast, none of these additional variables, with the only exception of one liquidity measure, can significantly predict excess returns in Eurodollar deposits.

While we do not study gold returns in dollar terms (or other currencies),[9] our results do

[9]The omission of gold prices from the analysis of gold lease rates has costs and benefits. On the one hand, we may lose useful information from the gold spot markets for the gold lending markets if such information is not included in a model of gold lease rates. However, Fama and French (1987) and Casassus and Collin-Dufresne (2005) show a weak link between gold prices and gold convenience yields and basis, so the lost information is likely to be small, if any. On the other hand, by not using gold spot markets, we avoid the potential short-run noises or stationary issues in gold prices, which can contaminate our inferences. We discuss this issue in more depth toward the end of this section.

suggest that the gold lease rates should be included as the "dividend" component of total gold returns, in addition to the "capital appreciation" component. For example, during the 16 years from June 1991 to June 2007, the annual appreciation of gold price is 3.56% on average, whereas the annual lease rate is 1.06% on average. In this period, gold lease rates contribute to 23% of total gold returns.

Our results provide new evidence on gold as an asset – from the perspective of gold borrowing and lending – and complement existing studies on gold *spot returns*. For example, Bailey and Chan (1993) find that a higher credit spread predicts a lower gold return, but we find the opposite for the excess return of gold bonds. Moreover, Barro and Misra (2013) document an insignificant covariance between gold spot returns and GDP growth, whereas we show that real GDP growth forecasts can predict excess returns of long-maturity gold bonds. Finally, while Erb and Harvey (2013) find that the dollar price of gold tends to increase as the US dollar depreciates, we show that the excess returns on gold bonds are positively predicted by both foreign currency returns and a measure of the dollar risk. While these results from monthly regressions have reduced power compared to our weekly analysis, they nonetheless suggest that gold leasing markets and gold lease rates serve as important avenues to better understand gold as an asset, as well as its linkage to other parts of financial markets and the broader economy.

The closest precursors to our analysis of gold lease rates, analogous to interest rates, are Schwartz (1997) and Casassus and Collin-Dufresne (2005). They model the joint dynamics of gold (and other commodities) convenience yields – conceptually similar to gold lease rates – together with gold spot prices and the instantaneous dollar interest rate. As shown by Joslin, Le, and Singleton (2012), however, the affine structure of their models necessarily forges a spanning relation between the commodity spot prices, the lease rates, and the dollar interest rates. The spanning condition implies that, for instance, the model-implied R-squared statistic from regressing the dollar gold prices onto the term structures of gold lease rates and dollar interest rates must be one. If this spanning relation does not hold in the data, forcing it to hold in the model will compromise the estimation of the commodity prices, the convenience yields, the interest rates, or any combination of the three.[10] This spanning issue does not arise in our stand-alone analysis of the gold lease rates because we do not jointly model the spot prices of gold or the dollar interest rates.

Instrumental to our analysis is the use of gold lease rates with *constant maturities*, implied

[10] Joslin, Le, and Singleton (2012) illustrate this point in the context of term structure models with spanned macroeconomic series, but the general spanning argument applies to many other affine models.

from gold forward rates. Compared to futures prices, the use of forward data avoids the potentially large measurement errors associated with delivery options in gold futures market. For example, investors shorting the June futures contract can choose any date in June to deliver. Thus, viewed by investors longing the June contract at the end of May, the effective maturity of the June contract could range from one to thirty days. Clearly, uncertainty about effective maturities leads to uncertainty about the implied yields, particularly for short maturities. This type of noise in futures data, combined with the relatively small magnitude of gold lease rates (about 1% in our sample), might have contributed to the view that gold has "insignificant" or "zero" convenience yield (see Schwartz 1997, for example). By contrast, we show that directly observed gold lease rates are economically significant and display high variability as well as significant time-varying risk premia.

2 Gold Lease Rates

2.1 Gold Lease Rates as Gold Interest Rates, Paid in Gold

Gold lease rates are "gold interest rate" investors pay in gold for borrowing gold. Although these rates are not directly quoted by banks, they can be constructed synthetically, by the no-arbitrage principle, as the difference between the unsecured U.S. dollar interest rate, denoted y, and the gold forward rate, denoted f, of the same maturity. In particular:

$$\text{Gold Lease Rate} = y - f, \tag{1}$$

where all rates are continuously compounding.

To see the logic behind (1), we consider the following simple example. Without loss of generality, suppose that today's gold price is S dollars per ounce and that we work with a one-year horizon. Today, an investor sells one ounce of gold and deposits the proceeds S in a bank account to earn Se^y dollars in one year. Simultaneously, the investor enters a forward contract today to buy back e^{y-f} ounces of gold in a year. Since the gold forward price is Se^f dollars per ounce, it costs the investor $e^{y-f} \times Se^f = Se^y$ dollars to accept delivery of the gold. This amount exactly offsets the investor's dollar balance in the US dollar deposit. Barring market frictions such as the bid-ask spread, this strategy is self-financing and gives the one-year expected return of $y - f$, in physical gold, on the investor's initial investment of one ounce of gold. Therefore, the gold lease rate, $y - f$, can be viewed as the "gold interest rate." Similar trading strategies can be implemented at other horizons to derive the term

structure of gold lease rates.

To proxy for the gold forward rate f, we use the "Gold Forward Offered Rate" (GOFO), adjusted to be continuously compounding. GOFOs are "rates at which dealers will lend gold on a swap basis against US dollars," according to the London Bullion Market Association (LBMA). In other words, GOFOs are the US dollar rates at which dealers are willing to pay in order to borrow US dollars, using physical gold as collateral.[11] At the maturity of the dollar loan collateralized by gold, the dollar borrower gets back his gold collateral and pays $X(1 + GOFO \times t)$ dollars, where X is the loan amount and t is the tenor of the loan. From the dollar borrower's perspective, the collateralized dollar loan can be decomposed into two trades: a spot sale of his gold at the price of X dollars and a forward repurchase of the same amount of gold at the price of $X(1 + GOFO \times t)$ dollars. Thus, GOFOs are also the gold forward rates measured in US dollars, just as the name implied. GOFOs are determined daily by the LBMA for 1-, 2-, 3-, 6- and 12-month maturities.[12]

To proxy for the U.S. dollar interest rate y, we use the Eurodollar deposit rates, also adjusted for the compounding frequency. The Eurodollar deposit rates are comparable to GOFOs since they cover similar maturities and both are borrowing rates for banks outside the U.S.

2.2 Data

The main data we use are the gold lease rates and Eurodollar deposits. GOFO data are obtained from LBMA, and Eurodollar deposit rates are obtained from Bloomberg. We construct the gold lease rate from the GOFO and Eurodollar data according to (1). All the spot rates are converted to a continuously-compounding basis. While LBMA's GOFO data start in July 1989, only since May 1990 do the GOFO data include all five maturities (i.e., 1-, 2-, 3-, 6- and 12-month). The Eurodollar deposit rates data with all five maturities are available starting June 1991. We exclude the financial crisis period in our sample. Thus, our data sample is from June 1991 to August 2007.

[11] Although the LBMA definition may appear to imply that the gold borrower pays GOFO, it is actually the lender of gold, namely the borrower of US dollars, who pays the interest rate.

[12] GOFOs are determined as follows. At 10.30 am London time, market makers in the London bullion markets enter their forward offered rates for all five maturities. A minimum of six contributors must enter rates in order for the means to be calculated. At 11.00 am, the mean is established for each maturity by discarding the highest and lowest quotations in each period and averaging the remaining rates. The current market making members of the LBMA include the Bank of Nova Scotia-ScotiaMocatta, Barclays Bank Plc, Deutsche Bank AG, HSBC Bank USA London Branch, Goldman Sachs, JP Morgan Chase Bank, Société Générale and UBS AG. See http://www.lbma.org.uk/pages/index.cfm?page_id=62 for more details.

To mitigate the effect of potential outliers in the data, we winsorize the data as follows. We first construct the five demeaned principal components (PCs) of the gold lease rates at the daily frequency. By construction, these five PCs are linear combinations of the five gold lease rates at different maturities. Next, we winsorize the PCs at the 1% level by their absolute values. Finally, we back out the implied gold lease rates from the winsorized PCs using the loading matrices in the original PC calculations. The same procedure is used on the Eurodollar deposit rates. Our analysis is conducted using these winsorized data at either the weekly frequency (for our main analysis in Section 3 and Section 4) or the monthly frequency (in Section 5). Winsorizing the data weakens the return predictability.

For our analysis in Section 5, we also include a range of other financial-market variables. The descriptions of those additional data are delegated to Section 5. In addition, we use the CME futures data as a robustness check, and the data description is delegated to Appendix A.

3 Expectation Hypothesis

In this section, we ask whether the expectation hypothesis (EH) can adequately characterize gold lease rates. We employ two most commonly used tests: (i) the Campbell and Shiller (1991) (CS) regression; and (ii) the Cochrane and Piazzesi (2005) (CP) regression.

3.1 The Campbell and Shiller Regression

We adapt the framework of CS and run the following the least-square regression:

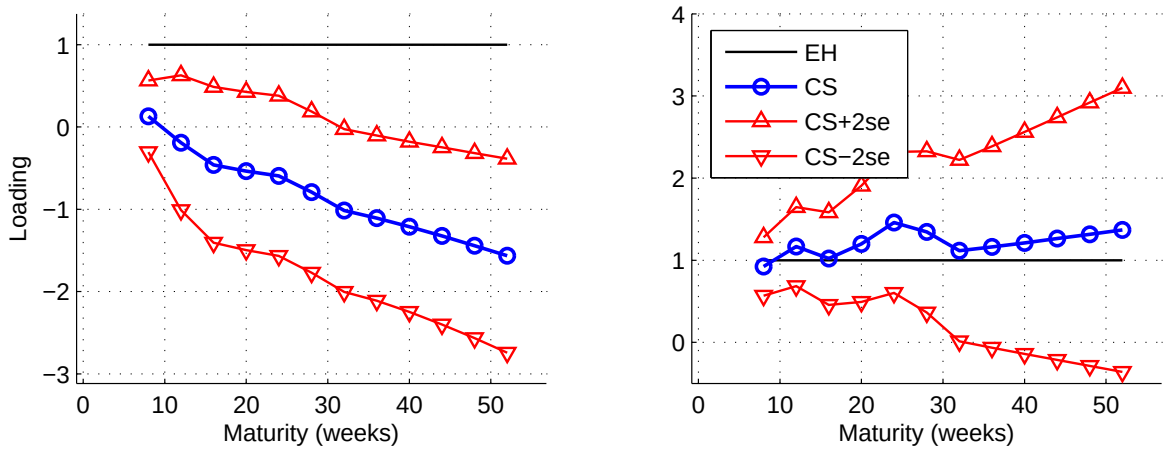
$$y_{t+4}^{n-4} - y_t^n = a_n + b_n \cdot \frac{4}{n-4} \cdot (y_t^n - r_t) + \epsilon_{t+4}^n, \tag{2}$$

where each period is one week, y_t^n denotes the n-week continuously-compounding yields at time t, and r_t denotes the 4-week short rate (y_t^4). All yield are annualized. The time t's are in increment of one week, and the maturity n's are in increments of four weeks from 8 weeks to 52 weeks. Under the EH, the coefficients b_n's should equal to one for all n's. Intuitively, if EH holds, the mark-to-market loss on long term loans should exactly offset the yield spread earned by the investor.[13]

[13] More rigorously, as explained by Dai and Singleton (2002), risk-premium adjusted expected changes in bond yields are proportional to the slope of the yield curve. In our setting, this means

$$E_t\left[y_{t+4}^{n-4} - y_t^n - (c_{t+4}^{n-4} - c_t^{n-4}) + \frac{p_t^{n-4}}{n-4}\right] = \frac{4}{n-4}(y_t^n - r_t),$$

Figure 2: CS Regression Coefficients.



(a) Gold Lease Rates

(b) Eurodollar Rates

Using risk-free interest rate data, CS show that the estimated b_n's are significantly below one and become increasingly negative for larger maturities n's. This pattern has been commonly considered among the most robust evidences against the EH. In applying this regression to the gold lease rates, we note that risk premia in gold lease rates may contain components for default and/or liquidity risks. In this sense, the null hypothesis we are testing is that the combined premia for all risks are constant. While we keep the notation "EH," we stress that our test is for this "generalized expectation hypothesis" (GEH), a term used in Downing and Oliner (2007).

Figure 2 (a) shows the CS coefficients for the gold lease rates together with their two-standard-error bounds. The standard errors are computed using the Newey and West (1987) covariance matrix with six lags. The CS regression loadings are significantly below one. The decreasing pattern of the loadings as maturity increases resembles what is observed with risk-free interest rates of much longer maturities. More strikingly, this pattern is in stark contrast to Figure 2 (b), where we plot the CS regression coefficients for the Eurodollar rates. For all maturities, the CS coefficients for the Eurodollar rates closely track the unit line, which is well within the two-standard-error bounds.

In the context of the Fed funds markets, results of Lange, Sack, and Whitesell (2003) suggest that patterns similar to Figure 2 (a) might be attributable to less transparent monetary

where c_t^n and p_t^n are the yield premium and forward premium, respectively. Under the EH, both c_t^n and p_t^n are constant for all n, predicting that $b_n = 1$ in the CS regressions.

policy in the early 90s, during which the Federal Reserve bank did not publicly announce policy changes and the rationale for their actions following the FOMC meetings. Indeed, Lange, Sack, and Whitesell (2003) find that after 1994, the fed funds rate became much more predictable and the EH performs much better. To check whether shifts in monetary policy over our sample, or more generally any structural break of similar nature, can explain the deviation from the EH observed for the gold lease rates, we split the our sample into sub-samples: before and after 1999. We choose the year 1999 because it is the year of the first central bank joint statement on gold (see Neuberger 2001).[14] The downward pattern of CS loadings found in Figure 2 (a) remains essentially intact for both sub-samples, suggesting that this pattern is unlikely sample specific.

In the context of the commercial paper markets, Downing and Oliner (2007) find evidence against the EH, but they also show that this violation of EH is entirely due to the rise in term premium around December, perhaps for "window dressing" reasons. Once this seasonal effect is controlled for, Downing and Oliner (2007) show that the EH actually holds in the data. To check whether our result is driven by a similar potential year-end effect, we drop all observations in the months of December and November from our sample, and find that Figure 2 (a) remains essentially the same. That is, the deviations from the EH observed with the gold lease rates cannot be explained by the year-end effects.

3.2 The Cochrane and Piazzesi Regression

Cochrane and Piazzesi (2005) regress bond excess returns on five annual forward rates (with maturities from one to five years) and find strong evidence of return predictability. Their results reveal that the inclusion of all yields, not only the slope, provides important additional information about risk premium. In this spirit, we implement the following regression:

$$rx_{t+4}^{n} = \alpha_n + \beta_n \mathcal{L}_t + \epsilon_{t+4}^{n}, \tag{3}$$

where rx_{t+4}^{n} denotes the 4-week excess return on an n-period bond, and where $\mathcal{L}_t$ denotes the vector of five zero yields, constructed from either the gold lease rates or Eurodollar rates, with maturities 4, 8, 12, 26, and 52 weeks.

Panels A and B of Table 1 report the results of this regression for the gold lease rates and the Eurodollar rates, respectively. The contrast between the predictability of gold bonds and Eurodollar excess returns remains remarkable. The adjusted R^2's across different maturities

[14] We also shorten our sample by starting from 1994 and our results remain robust.

in Panel A fluctuate around 20%, compared to just about 6% for Panel B. Additionally, the χ^2 tests of joint significance of the β coefficients are very strongly rejected for all regressions in Panel A, with probability values (pval) essentially at zeros. In contrast, the pval's of this test for two of four maturities in Panel B are either high or only borderline significant.

Table 1: Regression of 4-week excess returns onto yields with 4-, 8-, 12-, 26- and 52-week to maturity. $\overline{rx}$ denotes average excess returns across all maturities. Standard errors (bracketed) are computed using Newey West matrices with six lags. "χ^2 pval" refers to the probability value of the test of the joint significance of all the coefficient estimates. *, **, *** refer to the significance levels of 10%, 5%, and 1% respectively.

	4-week	8-week	12-week	26-week	52-week	χ^2 pval	Adj. R^2
Panel A: Gold Lease Rates							
$\overline{rx}$	-0.178^{**}	-0.281^{*}	0.562^{***}	-0.224^{*}	0.145^{**}	0.000	0.228
	(0.081)	(0.158)	(0.215)	(0.119)	(0.066)		
8-week	-0.069^{***}	0.041	0.071	-0.056^{*}	0.022	0.000	0.121
	(0.014)	(0.041)	(0.055)	(0.029)	(0.019)		
24-week	-0.143^{**}	-0.244^{*}	0.415^{**}	-0.030	0.025	0.000	0.178
	(0.059)	(0.140)	(0.192)	(0.099)	(0.060)		
40-week	-0.238^{**}	-0.381^{*}	0.759^{***}	-0.340^{**}	0.228^{***}	0.000	0.247
	(0.115)	(0.208)	(0.282)	(0.162)	(0.087)		
52-week	-0.295^{*}	-0.483^{*}	0.983^{***}	-0.501^{**}	0.325^{***}	0.000	0.254
	(0.161)	(0.256)	(0.342)	(0.209)	(0.105)		
Panel B: Eurodollar Rates							
$\overline{rx}$	0.093	0.154	-0.120	-0.299^{**}	0.183^{***}	0.022	0.059
	(0.102)	(0.321)	(0.332)	(0.126)	(0.067)		
8-week	-0.015	0.031	0.013	-0.050^{***}	0.022^{***}	0.043	0.060
	(0.015)	(0.035)	(0.029)	(0.017)	(0.007)		
24-week	0.048	0.079	-0.057	-0.140	0.076^{*}	0.268	0.034
	(0.086)	(0.233)	(0.251)	(0.089)	(0.045)		
40-week	0.147	0.243	-0.212	-0.441^{**}	0.279^{***}	0.016	0.065
	(0.144)	(0.458)	(0.475)	(0.185)	(0.100)		
52-week	0.234	0.380	-0.404	-0.536^{**}	0.350^{**}	0.026	0.063
	(0.194)	(0.602)	(0.623)	(0.269)	(0.147)		

4 No-arbitrage Term Structure Model of Gold Lease Rates

In this section, we lay down a standard discrete-time no-arbitrage term structure model of the gold lease rates. Our construction closely follows the canonical setup of Joslin, Singleton, and Zhu (2011) (JSZ), in which all normalizations for obtaining econometric identifications are implemented under the risk neutral measures. We then implement the full-information maximum-likelihood estimation and explore various implications of the model.

The use of no-arbitrage models offers several benefits, in addition to the regression-based analysis in Section 3. For one, a no-arbitrage model nests information across maturities in one coherent framework. Moreover, the model structure, in combination with Kalman filtering, allows for a robust treatment of potential measurement errors in the yield data. This is particularly important because measurement errors, left untreated, could result in potentially inflated fits for predictive regressions (see Cochrane and Piazzesi 2005 for a discussion of this point). Finally, the gold lease rates are zero-coupon rates. This is convenient for the estimation of no-arbitrage models because bootstrapping, typically required for coupon bond data, is unnecessary in our setting.

4.1 Modeling Setup

We let X_t be an N-element vector of state variables. The core of the risk-neutral setup consists of the risk neutral dynamics of states:

$$X_{t+1} = \lambda^{\mathbb{Q}} X_t + \sqrt{\Sigma_X}\epsilon^{\mathbb{Q}}_{t+1}, \tag{4}$$

with standard gaussian noise $\epsilon^{\mathbb{Q}}_{t+1}$ and the short rate specification:

$$r_t = r^{\mathbb{Q}}_{\infty} + \iota' X_t, \tag{5}$$

where r_t is the amount of gold (expressed as a percentage) earned on a gold bond over one unit of time interval; $\lambda^{\mathbb{Q}}$ is a diagonal matrix and, for simplicity, real-valued; and ι denotes a column vector of ones. It is important to stress that these normalizations, obtained through standard rotations of the state variables, are only for identification purposes.[15] Under the

[15]The assumption that $\lambda^{\mathbb{Q}}$ is real-valued is slightly over-identifying since it rules out the possibility that $\lambda^{\mathbb{Q}}$ may be complex. Nevertheless, as is shown in JSZ, the real-valued case is often empirically adequate.

risk neutral measures, investors are indifferent between making a long term loan and rolling short term loans successively. This indifference allows us to derive the n-period gold lease rates y_t^n for all n's as affine functions of the state variables:[16]

$$y_t^n = A_{n,X}(\lambda^{\mathbb{Q}}, r_\infty^{\mathbb{Q}}, \Sigma_X) + B_{n,X}(\lambda^{\mathbb{Q}})X_t. \tag{6}$$

We let y_t be the $J \times 1$ vector of yields used in estimation. As in JSZ, for each fixed $N \times J$ matrix W, we can write $\mathcal{P}_t = W\, y_t = W\, A_X + W\, B_X X_t$, where A_X is obtained by stacking $A_{n,X}$ on top of each other for all n's used in the estimation. B_X is similarly obtained by stacking column vectors $B_{n,X}$ together. This equation allows us to re-write the term structure model using $\mathcal{P}_t$ as the pricing factor:

$$y_t = A_{\mathcal{P}} + B_{\mathcal{P}}\mathcal{P}_t, \tag{7}$$

where $A_{\mathcal{P}} = A_X - B_{\mathcal{P}}(WA_X)$ and $B_{\mathcal{P}} = B_X(WB_X)^{-1}$.

The risk-neutral dynamics of $\mathcal{P}_t$ is given by:

$$\mathcal{P}_{t+1} = \kappa_0^{\mathbb{Q}} + \kappa_1^{\mathbb{Q}}\mathcal{P}_t + \sqrt{\Sigma_{\mathcal{P}}}\epsilon_{t+1}^{\mathbb{Q}}, \tag{8}$$

where $\Sigma_P = (WB_X)\Sigma_X(WB_X)'$, $\kappa_1^{\mathbb{Q}} = (WB_X)\lambda^{\mathbb{Q}}(WB_X)^{-1}$ and $\kappa_0^{\mathbb{Q}} = (I - \kappa_1^{\mathbb{Q}})WA_X$.

To complete the model, we specify the physical dynamics of $\mathcal{P}_t$ as:

$$\mathcal{P}_{t+1} = \kappa_0^{\mathbb{P}} + \kappa_1^{\mathbb{P}}\mathcal{P}_t + \sqrt{\Sigma_{\mathcal{P}}}\epsilon_{t+1}^{\mathbb{P}}, \tag{9}$$

where $\epsilon_{t+1}^{\mathbb{P}}$ follows a standard normal distribution, and $\kappa_0^{\mathbb{P}}$ and $\kappa_1^{\mathbb{P}}$ are both free parameters. As in JSZ, for each loading matrix W, the model is fully characterized by the parameter set $\Theta = \{\kappa_0^{\mathbb{P}}, \kappa_1^{\mathbb{P}}, r_\infty^{\mathbb{Q}}, \lambda^{\mathbb{Q}}, \Sigma_{\mathcal{P}}\}$.

4.2 Estimation Strategies

In fitting our model to the data, we employ two estimation strategies. The first approach is in the spirit of Chen and Scott (1993) in that exactly N combinations of yields are assumed to be observed perfectly. This allows the state factors to be conveniently inverted from the

[16] To be precise, the n-period gold lease rate y_t^n is per unit of time interval. That is, the gold price of a gold bond with n periods until maturity is given by $e^{-ny_t^n}$. The loadings can be computed as: $B_{n,X} = \frac{1}{n}diag((I - \lambda^{\mathbb{Q}})^{-1}(I - (\lambda^{\mathbb{Q}})^n))'$ and $A_{n,X} = r_\infty^{\mathbb{Q}} - \frac{1}{2n}tr\left(\Sigma_X \sum_{i=0}^{n-1} i^2 B_{i,X}B_{i,X}'\right)$.

observed yields portfolios. In the second approach, we allow all yields to be observed with errors and estimate the model through Kalman filtering.

The Chen and Scott approach

For the first strategy, we assume that the pricing portfolios, $\mathcal{P}_t$, are observed perfectly, whereas the remaining $J-N$ yields portfolios, $\mathcal{P}_t^e = W^e y_t$, are observed with i.i.d. errors. As is standard, W^e is rotated so that each of its rows, scaled to have unit L^2 norm, is orthogonal to each other and to each row of W. Following practices in the literature (see, for example, Joslin, Le, and Singleton 2012), we assume that the observation errors are uncorrelated and share one common variance. Specifically,

$$\mathcal{P}_t^{e,o} = \mathcal{P}_t^e + e_t \text{ with } e_t \sim N(0, \sigma_e^2 I_{J-N}), \tag{10}$$

where the superscript ^{o} differentiates the observed series from its theoretical counterpart.

The log likelihood of observing the entire time series of $(\mathcal{P}_t', \mathcal{P}_t^{e,o\prime})'$ is:

$$\mathcal{L} = \sum_t f(\mathcal{P}_{t+1}, \mathcal{P}_{t+1}^{e,o} | \mathcal{I}_t), \tag{11}$$

where f denotes the log conditional density and $\mathcal{I}_t$ denotes all information up to time t which is fully captured by $\mathcal{P}_t$.[17]

Maximum likelihood estimation requires the optimization of $\mathcal{L}$ over the parameter set Θ and the error variance parameter σ_e. As shown in JSZ, the global estimates of $\kappa_0^{\mathbb{P}}$ and $\kappa_1^{\mathbb{P}}$ are exactly the same as those obtained by fitting a VAR(1) process to $\mathcal{P}_t$. Additionally, in the Appendix, we show how we could concentrate out $(r_\infty^{\mathbb{Q}}, \sigma_e)$ analytically and $\Sigma_{\mathcal{P}}$ quasi-analytically. As such, the optimization is effectively only over the N diagonal elements of $\lambda^{\mathbb{Q}}$. Moreover, using the regression approach of Joslin and Le (2011), we obtain excellent

[17] $\mathcal{L}$ is computed as the sum of two components: $\sum_t f(\mathcal{P}_{t+1}|\mathcal{P}_t) + \sum_t f(\mathcal{P}_{t+1}^{e,o}|\mathcal{P}_{t+1})$. The first captures the time series dynamics and can be written as:

$$\sum_t f(\mathcal{P}_{t+1}|\mathcal{P}_t) = -\frac{T}{2} log((2\pi)^N |\Sigma_{\mathcal{P}}|) - \frac{1}{2}\sum_t ||\Sigma_P^{-1/2}(\mathcal{P}_{t+1} - (\kappa_0^{\mathbb{P}} + \kappa_1^{\mathbb{P}}\mathcal{P}_t)||_2^2,$$

where T refers to the sample length and $||.||_2$ denotes the L^2 norm. The second captures the cross-sectional fit and can be expressed as:

$$\sum_t f(\mathcal{P}_{t+1}^{e,o}|\mathcal{P}_{t+1}) = -\frac{T}{2} log((2\pi)^{J-N}\sigma_e^{2(J-N)}) - \frac{1}{2\sigma_e^2}\sum_t ||\mathcal{P}_{t+1}^{e,o} - W^e(A_{\mathcal{P}} + B_{\mathcal{P}}\mathcal{P}_{t+1})||_2^2.$$

starting guesses for $\lambda^{\mathbb{Q}}$. The end result is an instantaneous and very robust convergence for our estimation.

Estimation with filtering

For the second strategy, we relax the assumption that $\mathcal{P}_t$ is perfectly observed and assume that all yields are observed with i.i.d. errors whose joint distribution is non-degenerate. To set up the filtering problem, we use (9) as the state equation. For the observation equation, we again assume that the observational errors across maturities are uncorrelated and share one common variance:

$$y_t^o = A_{\mathcal{P}} + B_{\mathcal{P}}\mathcal{P}_t + e_t \text{ and } e_t \sim N(0, \sigma_e^2 I_J), \tag{12}$$

with $A_{\mathcal{P}}$ and $B_{\mathcal{P}}$ given in equation (7). Equipped with the state and observation equations, the likelihood of observing the entire panel of yields can be analytically computed through standard recursive procedures. Maximum likelihood estimation similarly requires the optimization of this likelihood over the same parameter set Θ and σ_e. Although the numerical optimization will be over the full parameter vector (since analytical concentration is no longer feasible), estimates from the first estimation approach can be used as the starting guesses. For our choices of the loading matrix W, these guesses are again excellent and deliver efficient and robust estimation.

Implementation

In implementing the estimation, we use the gold lease rates, for all available maturities (1-, 2-, 3-, 6-, and 12-month) from June 1991 to August 2007, constructed as described in Section 2.2. Notably, these rates correspond to straight gold bonds without coupons; hence no bootstrapping is required. We sample our data at the weekly frequency so the time interval for our model is $\Delta t = 1/52$. The weekly sampling of our data results in a time series with 848 data points.

Applying the principal component (PC) analysis to our gold lease rates data reveals that the first two PCs explain 99.7% of all variations, with the first PC explaining 96.7%. Given this evidence, we set the number of pricing factors to $N = 2$. Additionally, we choose the loading matrix W of the pricing portfolios to correspond to the those of the first two PCs.[18]

[18] As far as the first estimation approach is concerned, Joslin, Le, and Singleton (2012) show that this choice of the loading matrix is the most innocuous for it is likely to deliver estimates that are closest to those

For a comparable benchmark, we also fit our model to the Eurodollar dataset over the same sample, at the weekly frequency, for an identical maturities set, and with the same number of pricing factors ($N = 2$). We also choose the loading matrix W to correspond to the first two PCs of Eurodollar rates, although the actual numerical values of W will be slightly different between Eurodollar and the gold lease rates.

4.3 Parameter Estimates

Table 2 reports the estimates for $\kappa_1^{\mathbb{P}}$, $\kappa_1^{\mathbb{Q}}$, and the Cholesky decomposition of the covariance matrix $\Sigma_{\mathcal{P}}$ for the gold lease rates (Panel A) and the Eurodollar rates (Panel B). For each of the datasets, we report the estimates (Ests) together with corresponding robust standard errors (s.e.) using both the Chen and Scott (1993) approach (WOE) and the filtering approach (FY).[19]

For brevity, we have omitted the estimates of the intercept parameters ($\kappa_0^{\mathbb{P}}$, $r_\infty^{\mathbb{Q}}$) – these are available upon request. The pricing errors implied by the models are economically small with σ_e estimated at 7 basis points for the gold lease rates data (5 basis points for the Eurodollar data). These estimates are smaller than or comparable to estimates obtained in prior studies of the term structure of interest rates in the U.S. and other countries.

Two interesting observations emerge from the reported estimates. First, a quick comparison between the diagonal values of $\kappa_1^{\mathbb{Q}}$ and of $\kappa_1^{\mathbb{P}}$ suggests that the degree of persistence under $\mathbb{Q}$ is relatively higher than that under $\mathbb{P}$ for the gold interest rates, whereas the persistence in the Eurodollar data are similar under the $\mathbb{Q}$ and $\mathbb{P}$ measures. Explicit eigenvalue calculations of these matrices confirm this is indeed the case, for both the WOE and FY estimations. Since risk premia in no-arbitrage models are intimately connected to the difference between the physical and risk-neutral drifts, the difference in persistences under the two measures is informative of time-varying risk premia embedded in the gold lease rates – an issue we take on in depth in the next subsection. Second, while the estimates under WOE and FY are similar, they are not identical, perhaps with the exception of $\kappa_1^{\mathbb{Q}}$. The differences in

obtained by Kalman filtering. An important caveat here is Joslin, Le, and Singleton (2012)'s results are based on the interest rates data in the U.S., sampled at the monthly frequency. They also discuss conditions under which the filtering and the Chen and Scott (1993) approaches may be nearly equivalent. At the heart of these conditions is the relative difference between the magnitudes of time-series and cross-sectional errors, which can be quite different for distinct datasets and/or data frequencies.

[19] To compute these robust standard errors, we first express the first order conditions with respect to the parameter vector as relevant moment conditions in the Hansen (1982)'s GMM framework. We then use six lags in computing the Newey and West (1987) covariance matrix, which is prewhitened and recolored in accordance with Andrews and Monahan (1992) to treat possible autocorrelation in the residual errors.

Table 2: Parameter estimates (Ests) using the Chen and Scott approach (WOE), assuming the first two PCs of bond yields observed perfectly, and Kalman filtering (FY). Standard errors (s.e.) are computed using Newey West matrices with six lags. *, **, *** refer to the significance levels of 10%, 5%, and 1% respectively.

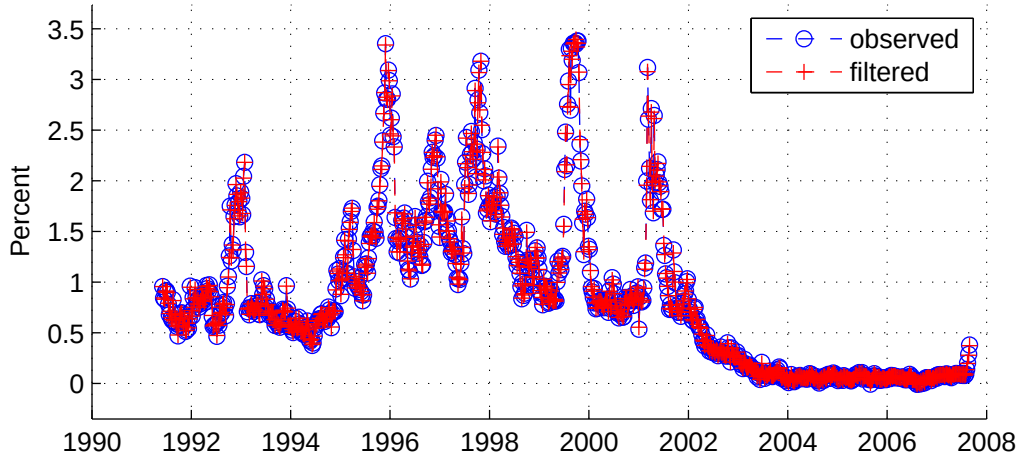
		$\kappa_1^{\mathbb{P}}$		$\kappa_1^{\mathbb{Q}}$		$\Sigma_{\mathcal{P}}$	
Panel A: Gold Lease Rate							
WOE	Ests	0.975***	0.026	0.994***	0.064***	0.167***	
		0.029**	0.864***	0.001	0.947***	−0.123***	0.179***
	s.e.	0.011	0.019	0.002	0.004	0.016	
		0.014	0.033	0.004	0.010	0.022	0.034
FY	Ests	0.978***	0.023	0.994***	0.061***	0.160***	
		0.028**	0.914***	0.001***	0.952***	−0.122***	0.117***
	s.e.	0.011	0.019	0.000	0.000	0.017	
		0.012	0.020	0.000	0.000	0.021	0.020
Panel B: Eurodollar Rate							
WOE	Ests	0.998***	0.058***	1.000***	0.051***	0.078***	
		−0.004*	0.967***	−0.005***	0.973***	0.016**	0.099***
	s.e.	0.002	0.011	0.000	0.002	0.005	
		0.002	0.010	0.001	0.006	0.007	0.007
FY	Ests	0.998***	0.058***	1.000***	0.051***	0.074***	
		−0.004**	0.980***	−0.005***	0.974***	0.017***	0.076***
	s.e.	0.002	0.011	0.000	0.002	0.005	
		0.002	0.009	0.001	0.006	0.006	0.006

estimates capture the differences between the observed yields portfolios $\mathcal{P}_t^o$ and their filtered counterpart. As we discuss shortly, the differences across WOE and FY estimates become more pronounced when these parameter estimates are transformed into those that govern the model's market prices of risks (reported in Table 3).

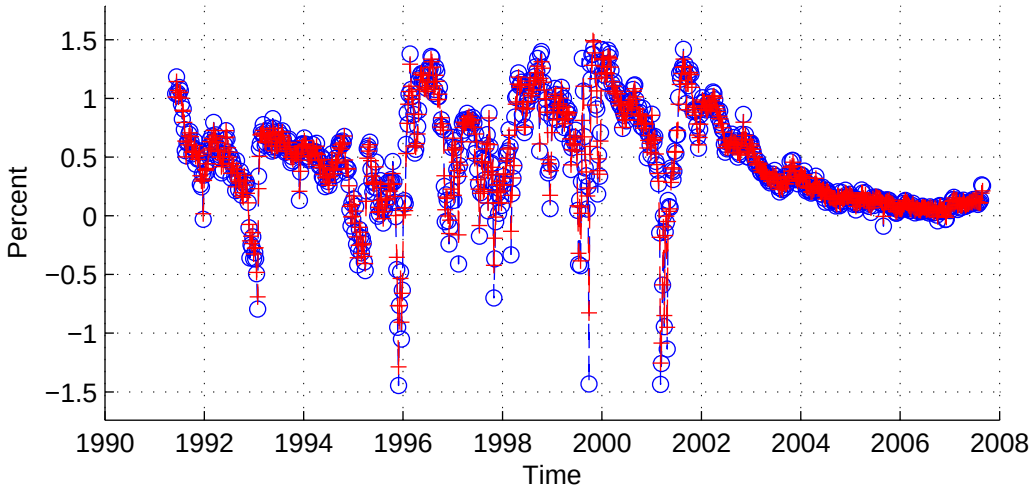
Figure 3 compares the observed series and its filtered counterpart of PC1 (subfigure (a)) and PC2 (subfigure (b)) of the gold lease rates. Similar to Joslin, Le, and Singleton (2012), we find that filtering is inconsequential for the level factor PC1 – the two graphs of subfigure (a) are on top of one another. Turning to the slope factor PC2, although the observed and filtered series track each other relatively well, filtering does have some effect in smoothing out some extreme movements of the slope. This suggests that observed PC2 may have nontrivial, albeit small, measurement errors.

Using the FY estimates, we plot in Figure 4 the expected yield curve (subfigure (a)) and

Figure 3: Observed and filtered time series of PC1 and PC2 of the gold lease rates.



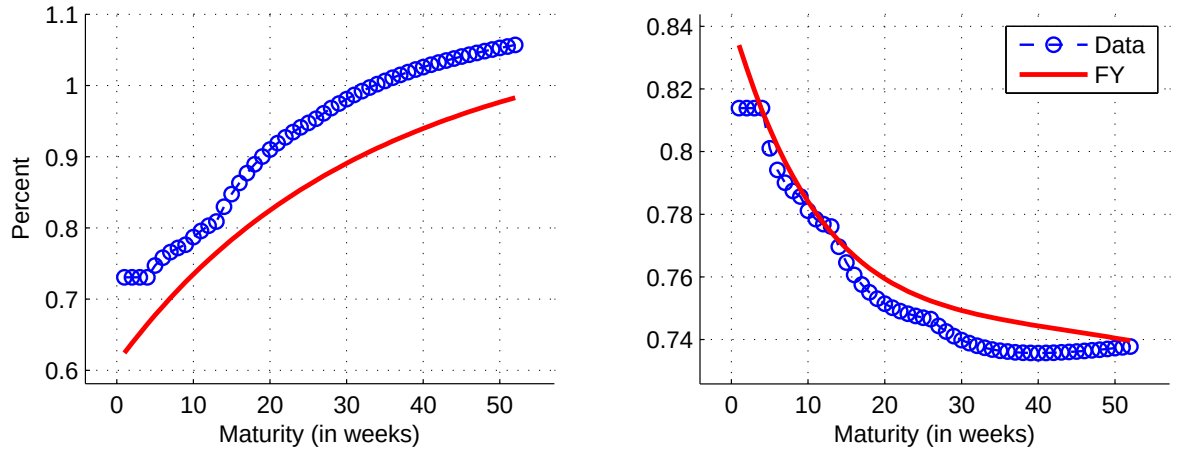
(a) PC1



(b) PC2

the term structure of rates volatility (subfigure (b)) implied by the no-arbitrage model against their sample counterparts for all maturities in one-week increment from one week to one year. Maturities not in our data (which contain only 1-, 2-, 3-, 6- and 12-month maturities) are filled by a standard bootstrapping technique that assumes constant forward rates between

Figure 4: Expected yields and term structure of yields volatility implied by the no-arbitrage model of gold lease rates. Estimation is done by Kalman filtering.



(a) Mean

(b) Volatility

available maturities (see, for example, Fama and Bliss 1987 for details).

The no-arbitrage model does a very good job in replicating both the upward pattern of the (zero-coupon) yield curve as well as the downward slope in rates volatility. Expected yield implied by the no-arbitrage model deviates from the corresponding sample averages by no more than ten basis points across all maturities. The upward-sloping pattern of the yield curve depends on the sign of the average risk premia implied by the model. We examine the model-implied dynamics of risk premia in detail in the next subsection.

The model-implied volatilities curve misses their empirical counterparts by no more than two basis points. The downward-sloping pattern of volatility is partly due to the stationarity of the states under the pricing ($\mathbb{Q}$) measures. If the eigenvalues of the risk-neutral feedback matrix $\kappa_1^{\mathbb{Q}}$ were outside of the unit circle, the yield loadings would be exploding as maturities increase, giving rise to an upward term structure of volatility. More subtly, the flexibility of the risk neutral setup, as opposed to the physical dynamics, allows our model to match the speed at which long-maturity yields converge to their limits (as maturities goes infinity), which in turn determines how fast or slowly volatility decreases in maturity.[20]

[20] Backus and Zin (1993) provides an example in which a no-arbitrage model, with constant market prices of risks, can grossly miss the volatility curve implied by the data. In their example, due to the constancy of the market prices of risks that shackles the physical and risk-neutral persistences together, their no-arbitrage model lacks the flexibility afforded by our setup.

4.4 Expectation Hypothesis and Time Varying Risk Premia

In this section, we consider the models' implication for the dynamics of risk premia and its consequence for the expectation hypothesis. At the heart of this analysis is the time-varying property of risk premia and their determinants.

We consider both variance risks (VAR) and standard deviation risks (STD). We let $\Sigma_{\mathcal{P}}^{1/2}$ be the Cholesky decomposition of the covariance matrix $\Sigma_{\mathcal{P}}$, and let $\mu_t^{\mathbb{P}}$ and $\mu_t^{\mathbb{Q}}$ be the physical and risk-neutral conditional means of the states $\mathcal{P}_t$, respectively:

$$\mu_t^{\mathbb{P}} = E_t^{\mathbb{P}}[\mathcal{P}_{t+1}] = \kappa_0^{\mathbb{P}} + \kappa_1^{\mathbb{P}}\mathcal{P}_t, \tag{13}$$

$$\mu_t^{\mathbb{Q}} = E_t^{\mathbb{Q}}[\mathcal{P}_{t+1}] = \kappa_0^{\mathbb{Q}} + \kappa_1^{\mathbb{Q}}\mathcal{P}_t. \tag{14}$$

The market prices of variance risks and standard-deviation risks can be computed, respectively, as:

$$\Lambda_t^{VAR} = \Sigma_{\mathcal{P}}^{-1}(\mu_t^{\mathbb{P}} - \mu_t^{\mathbb{Q}}) = \Sigma_{\mathcal{P}}^{-1}\left(\kappa_0^{\mathbb{P}} - \kappa_0^{\mathbb{Q}}\right) + \underbrace{\Sigma_{\mathcal{P}}^{-1}\left(\kappa_1^{\mathbb{P}} - \kappa_1^{\mathbb{Q}}\right)}_{\Lambda^{VAR}}\mathcal{P}_t, \tag{15}$$

$$\Lambda_t^{STD} = \Sigma_{\mathcal{P}}^{-1/2}(\mu_t^{\mathbb{P}} - \mu_t^{\mathbb{Q}}) = \Sigma_{\mathcal{P}}^{-1/2}\left(\kappa_0^{\mathbb{P}} - \kappa_0^{\mathbb{Q}}\right) + \underbrace{\Sigma_{\mathcal{P}}^{-1/2}\left(\kappa_1^{\mathbb{P}} - \kappa_1^{\mathbb{Q}}\right)}_{\Lambda^{STD}}\mathcal{P}_t. \tag{16}$$

Intuitively, (15) and (16) capture the extra expected returns required by investors for each extra "unit" in variance and standard deviation, respectively. In our setup, both (15) and (16) are affine in the state vector $\mathcal{P}_t$. The standard deviation risk formulation is more prevalent in existing term structure models, although the variance risk is easier to generalize (see, for example, Le, Singleton, and Dai 2010). The sensitivities of the two market prices of risk with respect to the state $\mathcal{P}_t$ are:

$$\Lambda^{VAR} \equiv \Sigma_{\mathcal{P}}^{-1}\left(\kappa_1^{\mathbb{P}} - \kappa_1^{\mathbb{Q}}\right), \tag{17}$$

$$\Lambda^{STD} \equiv \Sigma_{\mathcal{P}}^{-1/2}\left(\kappa_1^{\mathbb{P}} - \kappa_1^{\mathbb{Q}}\right). \tag{18}$$

Table 3 reports the market prices of risk parameters, Λ^{VAR} and Λ^{STD}, estimated from the no-arbitrage models for both the gold lease rates (Panel A) and the Eurodollar data (Panel B). Since we are primarily interested in the time-varying nature of the market prices of risks, we omit the intercept parameters in (15) and (16) in Table 3.

As we see from Panel A of Table 3, both entries in the second column of the market prices of risk matrix are significantly negative, whereas both entries in the first column are either

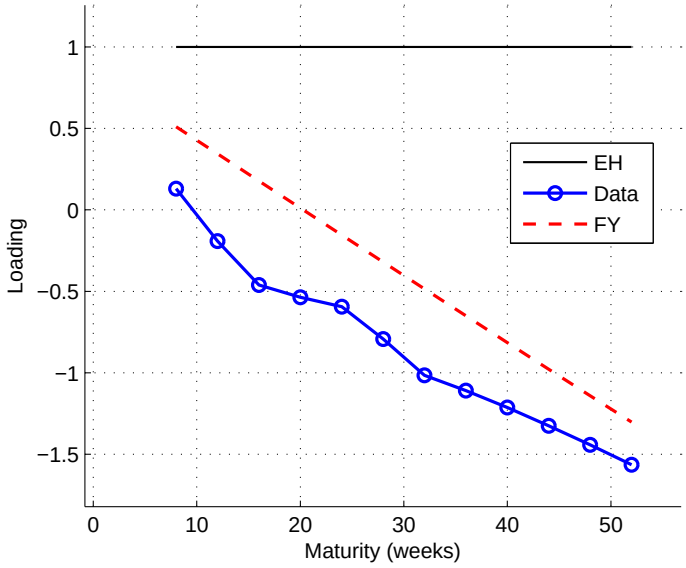
Table 3: Market prices of risks parameters estimates (Ests) using the Chen and Scott approach (WOE), assuming the first two PCs of bond yields observed perfectly, and Kalman filtering (FY). Standard errors (s.e.) are computed using Newey West matrices with six lags. *, **, *** refer to the significance levels of 10%, 5%, and 1% respectively.

		VAR risk		χ^2	STD risk		χ^2
Panel A: Gold Lease Rate							
WOE	Ests	-0.335	-3.853^{***}	0.000	-0.110^{*}	-0.223^{*}	0.000
		0.444	-3.431^{***}		0.080	-0.615^{***}	
	s.e.	0.458	0.868		0.064	0.122	
		0.368	0.662		0.070	0.099	
FY	Ests	0.175	-5.257^{***}	0.000	-0.102	-0.237^{*}	0.000
		1.062	-4.928^{***}		0.124	-0.574^{***}	
	s.e.	0.710	1.213		0.068	0.128	
		0.740	1.083		0.081	0.101	
Panel B: Eurodollar Rate							
WOE	Ests	-0.284	1.260	0.524	-0.021	0.087	0.529
		0.082	-0.685		0.008	-0.068	
	s.e.	0.245	1.508		0.020	0.124	
		0.211	1.008		0.021	0.100	
FY	Ests	-0.306	1.201	0.701	-0.020	0.103	0.693
		0.131	0.871		0.010	0.066	
	s.e.	0.284	1.897		0.021	0.143	
		0.339	1.665		0.026	0.125	

borderline or not significant. This is the case regardless of the estimation method, WOE or FY. In stark contrast, all of the entries estimated using the Eurodollar data, reported in Panel B, are individually insignificant, and they sometimes change signs across the two estimation methods. Moreover, when put together, estimates from Eurodollar data do not pass the χ^2 test of joint significance. As reported in the columns labeled χ^2, the pval's of the χ^2 test for Eurodollar data are always larger than 50%, irrespective of the estimation method (WOE or FY) or the type of risk (VAR or STD). By contrast, the pval's of the same tests for the gold interest rates, reported in Panel A, are essentially zeros.

The failure to reject the null hypothesis that the market prices of Eurodollar risks are constant is consistent with our earlier finding, reported in Figure 2 (b), that the Campbell and Shiller (1991) (CS) regression coefficients are indistinguishable from ones. On the flipped side, the strongly time varying nature of the market prices of gold lending risks suggests that

Figure 5: CS regression implied by the no-arbitrage model of gold lease rates. Estimation is done by Kalman filtering.



the corresponding CS regression coefficients will deviate from the unit line. Not immediately clear from our estimates, however, is the extent to which the model-implied CS regression coefficients match their sample counterparts.

Figure 5 plots the CS coefficients implied by the no-arbitrage model of gold lease rates, together with those implied by the data (which are also reported in Figure 2 (a)).[21] The model-implied regression coefficients (red crosses) clearly stay closer to the empirical counterparts (green squares) than to the unit line implied by the EH. Although the red line lies strictly above the green line for all maturities, based on evidences from existing studies using U.S. interest rates data (for example Dai and Singleton 2002), our model does a relatively good job in capturing the predictability implicit in the CS regressions. The goodness-of-fit of our model is comparable to previously estimated Gaussian term structure models of interest rates.

Given the significant market prices of risk parameters that are estimated from the gold lease rates data, a natural question entails: what do those estimates capture economically? To answer this question, it is instructive to examine the exact connection between the market prices of risk matrix and the model-implied risk premia (or, equivalently, the model-implied

[21] Dai and Singleton (2002) provide the analytical expressions to compute the CS coefficients from model parameters.

expected excess returns). For example, consider a generic one-period security whose payoff next period is $e^{-c'\mathcal{P}_{t+1}}$, where c is a constant vector that determines the exposure of this security to the factor risk in next period's state variable $\mathcal{P}_{t+1}$. The continuously compounded expected excess return on the security is given by:

$$\begin{aligned} E_t^{\mathbb{P}}\left[\log\left(\frac{e^{-c'\mathcal{P}_{t+1}}}{E_t^{\mathbb{Q}}[e^{-r_t}e^{-c'\mathcal{P}_{t+1}}]}\right)\right] - r_t &= -c'(\mu_t^{\mathbb{P}} - \mu_t^{\mathbb{Q}}) + \text{constant}, \\ &= -c'\Sigma_{\mathcal{P}}\Lambda^{VAR}\mathcal{P}_t + \text{constant}. \end{aligned} \tag{19}$$

That is, the time-varying part of the expected excess return for holding this security is equal to the quantity of variance risks, $-c'\Sigma_{\mathcal{P}}$, multiplied by the market prices of variance risks, $\Lambda^{VAR}\mathcal{P}_t$.

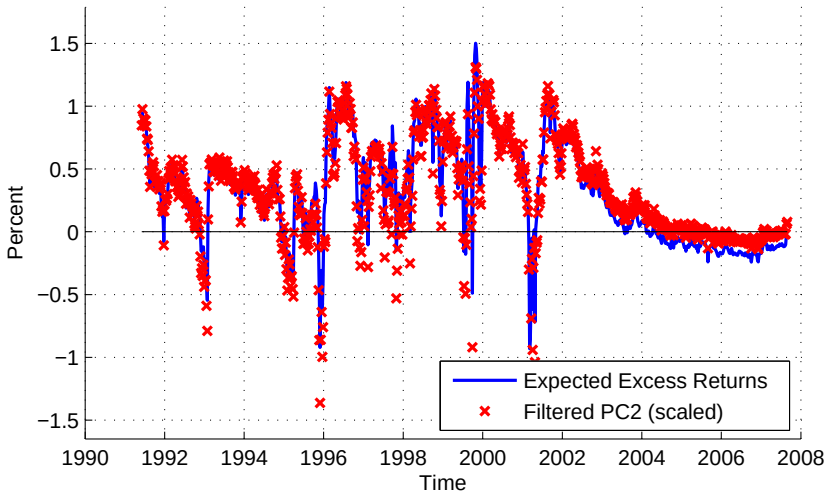
To further illustrate the economic intuition, consider now a security whose risk exposure is such that $c'\Sigma_{\mathcal{P}} = (1, 0)$, i.e., the only relevant source of risk is the variance risk of the level factor. For such a security, the time variation in risk premia is driven by the product of the first row of Λ^{VAR} and the state vector $\mathcal{P}_t$. Combined with the fact that only the second column of Λ^{VAR} is significantly estimated (see Table 3), it implies that the slope factor, or PC2, is virtually the only determinant of premia that compensates the level risk. Likewise, for a slope-risk-only security (i.e. $c'\Sigma_{\mathcal{P}} = (0, 1)$), slope is essentially the only factor relevant for risk premia because the (2,1) entry of Λ^{VAR}, which picks up the level factor's effect on risk premia, is insignificant.

The negative signs for the (1,2) and (2,2) entries of Λ^{VAR}, after interacting with the negative sign accompanying yields when converted to bond prices, translate into a positive relationship between slope and expected excess returns. Holding all else constant, an increase in the steepness of the gold interest rates curve leads to an increase in the required compensation for bearing gold lending (variance) risks.

Our line of reasoning also applies to standard deviation risks. In short, the negative entries observed for the second column of Λ^{STD} imply that the standard deviation risks of both level and slope factors are priced. The only factor that determines the compensation for these risks is the steepness of the gold interest rates curve. The steeper the curve, the higher is the risk premium required for taking on gold lending (standard deviation) risks.

To visually confirm the linkage between slope and expected excess returns, we perform the following exercise. We take each maturity used in the estimation (1-, 2-, 3-, 6-, and 12-month),

Figure 6: Expected weekly-excess returns (annualized) and filtered slope.



compute the annualized one-week expected excess returns implied by the FY model,[22] and plot the cross-maturity averages in Figure 6. Next, we plot on the same graph the filtered time series of the slope (PC2), which is normalized to have the same mean and standard deviation as the first series. Evidently, the two series bear remarkable resemblances with a correlation of 0.99. On the other hand, the correlation between level and the model-implied expected excess returns is essentially zero.

It is evident from Figure 6 that risk premia in gold leasing markets are mostly positive, with the exception of the last three years of the sample. This explains why the average term structure of gold lease rates is upward sloping, as is seen in Figure 4 (a).[23] The average positivity of risk premia also suggests that buying a gold bond (or lending gold) is often viewed by the marginal investor as a risky investment. For this to be the case, upward shifts

[22]For a gold bond with n weeks to maturity, the one-week expected excess return can be computed as

$$nA_{\mathcal{P},n} - (n-1)A_{\mathcal{P},n-1} - A_{\mathcal{P},1} - (n-1)B_{\mathcal{P},n-1}\kappa_0^{\mathbb{P}} + \left(nB_{\mathcal{P},n} - (n-1)B_{\mathcal{P},n-1}\kappa_1^{\mathbb{P}} - B_{cP,1}\right)\mathcal{P}_t,$$

where $A_{\mathcal{P},n}$ and $B_{\mathcal{P},n}$ are the intercept and loadings for the n-week maturity.

[23]For a formal derivation, let us denote the one-week excess return on the n-period gold bond by

$$xr_t^n = ny_t^n - (n-1)y_{t+1}^{n-1} - r_t,$$

and note that

$$\frac{1}{n}\sum_{i=0}^{n-1} xr_{t+i}^{n-i} = y_t^n - \frac{1}{n}\sum_{i=0}^{n-1} r_{t+i},$$

in the gold yield curve (resulting in reduction in bond prices) must on average be perceived as bad news.

Equally notable from Figure 6 are the occasional short episodes during which risk premia become negative. Thus once in a while buying a gold bond can provide some insurance benefits. For example, the cluster of negative risk premia observed during December 1994 and for much of 1995 coincides with Orange County's bankruptcy, which, according to the New York Times on December the 8th 1994, "shook the market for public borrowing across the country." It is perhaps possible that under such extraordinary circumstances, investment in gold bonds provide insurance-like benefits for holders of gold, just like a US treasury bond is a safe haven for keeping US dollars in times of crisis. It is important to note, however, that unlike treasury bonds, gold bonds are typically not free of default risks. As a result, Figure 6 suggests that the crisis-insurance benefits of gold bonds can more than offset default risk, even during market stress.

4.5 Sub-sample Analysis

In this subsection we conduct a robustness check of our main result. Neuberger (2001) suggests that the 1999 joint central bank statement on gold may be pivotal for the predictability embedded in gold interest rates. As a robustness check, we divide our sample into two sub-samples, June 1991–August 1999 and October 1999–August 2007. We leave out September 1999, the monthly in which the first joint central bank statement on gold was issued.

For both of the sub-samples, the market price of standard deviation risks parameters continue to be jointly significant for the gold lease rates, as we see from the joint χ^2 statistics of Table 4. The slope is again the dominant determinant of risk premia. Similar to the full-sample results, the slope loadings (the second column of Λ^{STD}) are, except for one entry, significantly negative with broadly comparable magnitudes. The (borderline) non-significance of the (1,2) entry of Λ^{STD} in the first sub-sample is at least partly due to the weaker statistical power given the shorter sample. Notably, in the second sub-sample, there is some evidence that the level factor, in addition to slope, may play a role in determining premia for

which, by applying the expectation operator to both sides, yields:

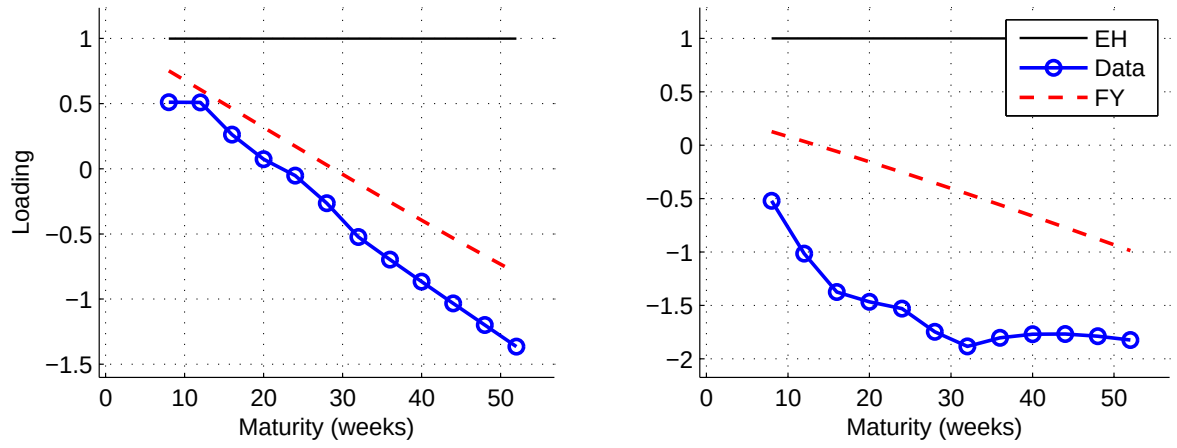
$$\frac{1}{n}\sum_{i=0}^{n-1} E[xr_{t+i}^{n-i}] = E[y_t^n - r_t].$$

Clearly, the average slope, as measured by $E[y_t^n - r_t]$, must be positive, implying an upward-sloping average yield curve, so long as the risk premia are also positive on average.

Table 4: Sub-sample Analysis using gold lease rates. Market prices of risks parameters estimates (Ests) using the Chen and Scott approach (WOE), assuming the first two PCs of bond yields observed perfectly, and Kalman filtering (FY). Standard errors (s.e.) are computed using Newey West matrices with six lags. *, **, *** refer to the significance levels of 10%, 5%, and 1% respectively.

		June 1991 - Aug 1999			Oct 1999 - Aug 2007		
		STD risk		χ^2	STD risk		χ^2
WOE	Ests	-0.110	-0.223	0.000	-0.110^{**}	-0.223^{**}	0.005
		0.080	-0.615^{***}		0.080	-0.615^{*}	
	s.e.	0.096	0.121		0.183	0.213	
		0.127	0.105		0.208	0.243	
FY	Ests	-0.102	-0.237	0.006	-0.102^{*}	-0.237^{**}	0.002
		0.124	-0.574^{***}		0.124	-0.574	
	s.e.	0.109	0.134		0.220	0.211	
		0.159	0.120		0.245	0.266	

Figure 7: Sub-sample analysis. CS regression implied by the no-arbitrage model of gold lease rates. Estimation is done by Kalman filtering.



(a) Jun 1991 - Aug 1999

(b) Oct 1999 - Aug 2007

compensating level risks.

Figure 7 reports the CS regression coefficients which, for both sub-samples, again lie significantly below the unit line. Importantly, the no-arbitrage models estimated for each

sub-sample are again able to replicate to a reasonable extent these deviations from the expectation hypothesis.

We find that the different behavior of risk premia in gold lease rates and Eurodollar rates persist in each sub-sample. For Eurodollar rates, the estimates of Λ^{STD} continue to be jointly insignificant in each sub-sample; so are results using Λ^{VAR}. The CS coefficients using the Eurodollar data are again indistinguishable from the unit line for each sub-sample. These results using Eurodollar data are omitted for brevity and are available upon request.

5 Financial Market Conditions and Gold Lease Rates

Evidence from Section 3 and Section 4 demonstrates that expected excess returns of "gold bonds" – lending gold for a long maturity and borrowing gold for a short maturity – are strongly predictable, especially by the slope of the gold lease rates. We also find sharply different behaviors of risk premia in gold lease rates versus Eurodollar deposit rates at the same maturities. A natural question is why. In this section, we analyze this question by exploring the effects of a variety of additional financial-market factors for risk premia in the two markets. We further show that: (i) the predictive power of the slope factor of gold lease rates is robust (even stronger) after including other variables; (ii) several additional variables predict excess returns in gold lease rates above and beyond the PCs; and (iii) Eurodollar PCs and the additional variables have little predictive power for risk premia in Eurodollar markets.

We complement the data on gold lease rates and Eurodollar deposits by the following variables.[24] Because the macro variables are only available at the monthly frequency, all variables below are monthly.

- Forecasts of inflation and real GDP growth. Because realized macroeconomic activities are reported with a lag, we use forecasts of inflation and real GDP growth of the United States, obtained from the Bluechip survey newsletters. Each month, Bluechip forecasts provide survey-based annualized inflation estimates and real GDP estimate for the next 1, 2, 3, 4, 5 and 6 quarters.[25] Given the noise in short-term inflation and GDP forecasts, we calculate the expected inflation (Einf) and expected real GDP growth (Ergdp) as the average of the corresponding 3-, 4-, 5- and 6-quarter forecasts.

[24] We thank the authors mentioned below for making their data available.

[25] Ang, Bekaert, and Wei (2007) find that survey data provide better inflation forecasts than macro variables and asset markets do.

- Gold returns and volatility. We calculate the monthly realized returns (goldRet) and volatilities (goldVol) of gold from daily London PM gold fixing, provided by LBMA and downloaded from Bloomberg. The gold return volatilities are annualized.

- Equity market factors. We include the dividend yield of value-weighted NYSE/Nasdaq/Amex stocks (divYield), downloaded from CRSP, and the Fama and French (1992) three factors (the market excess return equityVW, SMB, and HML), downloaded from Ken French's website.

- Credit market factor. We use the Baa-Aaa credit spread (Baa-Aaa), provided by Moody's and downloaded from Bloomberg.

- Currency market factors. We include the average foreign currency excess returns (FXRet) and the carry-trade factor (FXHML) of Lustig, Roussanov, and Verdelhan (2011). We also include the "global component of the dollar factor" (dollarGlobal) of Verdelhan (2013). These three variables are downloaded from Adrien Verdelhan's website.

- Liquidity factors. We include the bond liquidity measures by Fontaine and Garcia (2012) (FG), Wright and Zhou (2009) (WZ), and Hu, Pan, and Wang (2012) (HPW). The FG and WZ measures are obtained from the authors, and the HPW measure is downloaded from Jun Pan's website. Because these variables have different magnitudes, we standardize them, so that they have means of zero and standard deviations of one.

- Volatility factors. We include three variables. The first is the VIX index, downloaded from CME's website. The second is the variance risk premium of Bollerslev, Tauchen, and Zhou (2009) (VRP), defined as the different between the implied stock-market variance over the coming month and the realized stock-market variance over the past month; it is downloaded from Hao Zhou's website. The third is the Merrill Option Volatility Estimate (MOVE) index, a proxy for treasury bond yield volatility; it is downloaded from Bloomberg. The volatility factors are also standardized, so that they have means of zero and standard deviations of one.

Table 5 summarizes the mean and standard deviations of the first two PCs of gold lease rates and Eurodollar rates, as well as the mean and standard deviations of the additional variables. Note that the PCs are demeaned already. The liquidity and volatility factors are already standardized, so they are omitted from the table. The final dataset has 195 months

from June 1991 to August 2007. We denote by X_t the vector of the above additional variables observed at month t.

Table 5: Summary statistics of monthly data.

	Gold PC1	Gold PC2	Eurodollar PC1	Eurodollar PC2	Einf	Ergdp	goldRet	goldVol
Mean	0	0	0	0	0.0271	0.0293	0.0032	0.1184
St.Dev.	0.0165	0.0033	0.0368	0.0030	0.0048	0.0050	0.0366	0.0597
	divYield	equityVW	SMB	LMH	Baa-Aaa	FXRet	FXHML	dollarGlobal
Mean	0.0016	0.0063	0.0020	0.0046	0.0082	0.0017	0.0085	-0.0024
St.Dev.	0.0006	0.0402	0.0369	0.0340	0.0019	0.0157	0.0254	0.0282

As in Section 3, we denote by y_t^n the period-t spot rate of a n-period gold bond and denote by r_t the period-t spot rate of a 1-period gold bond. (In this section one period is one month.) The excess return of holding an n-period gold bond for one month is

$$rx_{t+1}^n \equiv y_t^n \frac{n}{12} - y_{t+1}^{n-1} \frac{n-1}{12} - r_t \frac{1}{12} = \frac{y_t^n - r_t}{12} + (y_t^n - y_{t+1}^{n-1}) \frac{n-1}{12}. \tag{20}$$

This excess return is measured in gold. We run the predictive regression for the average excess return across maturities $n \in \{2, 3, 6, 12\}$:

$$\overline{rx}_{t+1} \equiv \frac{1}{4} \sum_n rx_{t+1}^n = \overline{\alpha} + \overline{\beta} \cdot [goldPC, EDPC, X]_t + \overline{\epsilon}_{t+1}, \tag{21}$$

where $goldPC$ and $EDPC$ are the first two PCs of the gold lease rates and Eurodollar deposit rates, respectively. We also run the same regression for excess returns for individual maturities:

$$rx_{t+1}^n = \alpha_n + \beta_n \cdot [goldPC, EDPC, X]_t + \epsilon_{t+1}^n. \tag{22}$$

Table 6 reports the results of regression (21) applied to average monthly excess returns for gold lease rates and Eurodollar deposit rates. For the ease of exposition, the reported coefficients in the tables are multiplied by 100. These regressions reveal three salient features of the excess returns.

First, regressions with only PCs as predictors (i.e. the first column of the two regressions) confirm our earlier evidence for the rejection of the expectation hypothesis using yield-only models. The PCs of gold lease rate significantly predict excess returns of gold bonds. PC1 and PC2 explain 10% of the average excess returns. The inclusion of the other financial-market variables increases the coefficient on gold lease rate PC1 by a factor of about 6, and PC2,

by a factor about 2. By contrast, average excess returns on Eurodollar deposits cannot be explained by PC1 and PC2 of Eurodollar rates at all. The additional variables make Eurodollar PC1 significant at 10% level, but does not change the insignificance of Eurodollar PC2.

Second, several financial-market variables have incremental explanatory power for excess returns in gold lease rates, as we see in the second column of Table 6. They increase the adjusted R^2 from 10.0% to 22.7%. The first two PCs of the Eurodollar deposit rates are not significant. Between inflation forecast and real GDP growth forecast, only the latter is significant (at the 10% level). A 1% (or about two-standard deviation) increase in the real GDP growth forecast raises the average monthly excess return of gold bonds by 7.8 bps per month, or about 94 bps per year. That is, excess returns on gold bonds move almost one-for-one with expected real GDP growth. Expected inflation is not significant. If gold bond risk premia at least partly capture the expected real returns in the economy, we would naturally expect a positive coefficient on expected real GDP growth but insignificant coefficient on expected inflation.

Between the two gold-market factors, gold spot returns have no predictive power, but gold volatility positively predicts excess returns on gold bonds. The effect of volatility is natural because gold lenders, who may also care about the dollar value of gold, would demand a higher risk premium in the presence of a higher gold volatility.

Among factors in other asset classes, the small-firm premium (SMB), credit spread (Baa-Aaa), foreign currency returns (FXRet), and the global component of dollar risk (dollarGlobal) are all significant. For example, the ex ante measure of credit risk premium, the Baa-Aaa credit spread, positively predicts excess returns of gold bonds. Since Bailey and Chan (1993) find that the credit spread predicts gold *spot* returns with a negative sign, our evidence suggests that gold lease rates contain a risk premium that behaves different from that in physical gold returns. Moreover, since gold is a global asset, we would expect it to pick up some global risk factors. Indeed, global risks reflected in the foreign currencies (FXRet) and the US dollar (dollarGlobal) can predict excess returns on gold bonds with positive signs. On the other hand, the sign on SMB is negative. It indicates that risk premia associated with small firms and risk premia in credit markets and currency markets behave differently.

None of the six liquidity and volatility factors are significant. Since those factors are from the equity markets and bond markets, the lack of significance suggests that risk premia in gold lease rates do not reflect liquidity conditions in those markets.

Third, for the Eurodollar deposits, even though the additional variables increase the

adjusted R^2 of the predictive regression from 0% to 20.4%, these variables are not significant, with the only exception of the WZ liquidity measure. The negative sign on the WZ measure is consistent with the original Wright and Zhou (2009) result. In particular, factors that explain excess returns on gold bonds are not significant for Eurodollar deposit rates. This evidence, again, points to different behaviors of risk premia in the two markets.

Table 7 and Table 8 report results from regression (22), using individual maturities. For the gold lease rates (Table 7), factors that are significant for average-return regression – the first two gold lease rate PCs, gold volatility, SMB, foreign currency returns, and the global risk of the US dollar – are also consistently significant in the individual-maturity regressions. Moreover, the magnitudes of their coefficients increase in maturity, as we have expected. The expected real GDP growth (Ergdp) only has significance in the 12-month regression, but its magnitude there is much larger. This is not too surprising since the expected real GDP growth is calculated by averaging the 3- to 6-quarter ahead forecasts. The HPW "noise" measure also becomes significant in the 12-month regression.

In contrast, the individual-maturity regressions for Eurodollar deposits in Table 8 show much weaker predictability, with the WZ liquidity measure being the only consistently significant predictor. PC1 of Eurodollar deposit becomes significant only for the 12-month regression. For the 3-month and 6-month regressions only, dividend yield (divYield) and HPW noise measure become positively significant. Interestingly, for the 2-month regression, expected inflation becomes significant at the 10% level with a positive sign, which is consistent with the intuition that a higher expected inflation demands a higher risk premium on nominal deposits. These regressions, again, confirm out results in the average-return regression that risk premia in gold lease rates are strongly predictable, whereas the risk premia in Eurodollar deposits are much less so.

Table 6: Average excess return regression (21) applied to gold lease rates and Eurodollar deposit. The left-hand side is the average monthly excess return of holding a gold bond or a Eurodollar deposit with 2-, 3-, 6-, and 12-month maturities. The PCs are normalized to have mean zero. Variables FG, HPW, WZ, VIX, VRP, and MOVE are normalized to have mean zero and standard deviation one. All the regression coefficients are multiplied by 100. The t-statistics are reported in parentheses, and statistical significance at the 10%, 5%, and 1% levels are denoted by *, **, and ***, respectively.

	Gold lease rates		Eurodollar deposit rates	
Const	0.027^{***}	-0.449^{***}	0.024^{***}	-0.153
	(3.907)	(−2.930)	(4.155)	(−1.524)
goldPC1	0.700	4.062^{***}		0.421
	(1.125)	(3.341)		(0.696)
goldPC2	9.831^{***}	17.654^{***}		0.485
	(3.413)	(4.545)		(0.217)
EDPC1		−0.327	0.164	0.454^{*}
		(−1.142)	(1.237)	(1.725)
EDPC2		2.091	−0.081	0.719
		(0.606)	(−0.032)	(0.253)
Einf		2.987		3.234
		(1.223)		(1.411)
Ergdp		7.764^{*}		0.929
		(1.914)		(0.447)
goldRet		0.077		−0.033
		(0.409)		(−0.236)
goldVol		0.552^{***}		−0.127
		(4.088)		(−1.323)
DivYield		5.562		15.437
		(0.436)		(1.210)
EquityVW		0.267		−0.187
		(1.141)		(−1.034)
SMB		-0.555^{**}		−0.023
		(−2.443)		(−0.157)
HML		0.054		0.071
		(0.200)		(0.315)
Baa-Aaa		11.319^{***}		6.346
		(2.700)		(1.546)
FXRet		1.401^{**}		0.437
		(2.001)		(0.621)
FXHML		−0.151		0.053
		(−0.506)		(0.218)
dollarGlobal		0.729^{*}		−0.142
		(1.952)		(−0.385)
FG		0.013		−0.003
		(1.250)		(−0.309)
HPW		−0.019		0.008
		(−1.560)		(0.903)
WZ		0.002		-0.034^{***}
		(0.168)		(−4.316)
VIX		−0.019		−0.011
		(−1.144)		(−0.925)
VRP		0.013		0.010
		(1.255)		(1.353)
MOVE		−0.008		−0.010
		(−1.119)		(−1.632)
R2 adj	0.100	0.227	−0.005	0.204

Table 7: Excess return regression (22) applied to gold lease rates. The left-hand side is the monthly excess return of holding a 2-, 3-, 6-, or 12-month gold bond for 1 month. The PCs are normalized to have mean zero. Variables FG, HPW, WZ, VIX, VRP, and MOVE are normalized to have mean zero and standard deviation one. All the regression coefficients are multiplied by 100. The t-statistics are reported in parentheses, and statistical significance at the 10%, 5%, and 1% levels are denoted by *, **, and ***, respectively.

	2-month gold bond		3-month gold bond		6-month gold bond		12-month gold bond	
Const	0.008^{***}	-0.101^{**}	0.013^{***}	-0.126	0.040^{***}	-0.441^{**}	0.047^{***}	-1.127^{***}
	(3.215)	(−1.974)	(3.052)	(−1.303)	(4.897)	(−2.468)	(3.288)	(−3.464)
goldPC1	0.606^{**}	1.434^{***}	0.763^{*}	2.099^{**}	0.848	4.552^{***}	0.585	8.164^{***}
	(2.112)	(2.874)	(1.750)	(2.312)	(1.236)	(3.175)	(0.456)	(3.374)
goldPC2	−0.580	1.135	1.829	4.131^{*}	10.866^{***}	19.701^{***}	27.210^{***}	45.649^{***}
	(−0.431)	(0.756)	(1.057)	(1.653)	(3.061)	(4.272)	(4.833)	(5.729)
EDPC1		−0.080		−0.197		−0.425		−0.605
		(−0.791)		(−1.080)		(−1.280)		(−1.008)
EDPC2		0.888		2.227		0.542		4.709
		(0.721)		(0.988)		(0.134)		(0.670)
Einf		0.355		0.340		2.646		8.605
		(0.398)		(0.243)		(0.952)		(1.620)
Ergdp		1.977		1.467		8.267		19.346^{**}
		(1.318)		(0.486)		(1.639)		(2.508)
goldRet		0.065		0.046		0.142		0.055
		(0.849)		(0.370)		(0.677)		(0.133)
goldVol		0.171^{***}		0.317^{***}		0.576^{***}		1.145^{***}
		(3.472)		(3.943)		(3.940)		(3.722)
DivYield		2.309		5.441		2.731		11.769
		(0.513)		(0.672)		(0.180)		(0.448)
EquityVW		0.077		0.124		0.140		0.729
		(0.872)		(0.747)		(0.515)		(1.522)
SMB		-0.202^{**}		-0.257^{*}		-0.762^{***}		-0.998^{**}
		(−2.454)		(−1.800)		(−2.837)		(−2.090)
HML		−0.074		−0.032		−0.196		0.516
		(−0.709)		(−0.157)		(−0.568)		(1.053)
Baa-Aaa		2.314		4.994^{*}		11.807^{**}		26.161^{***}
		(1.436)		(1.866)		(2.392)		(2.929)
FXRet		0.605^{*}		1.199^{**}		1.661^{**}		2.138
		(1.962)		(2.487)		(2.051)		(1.527)
FXHML		−0.138		−0.084		−0.258		−0.125
		(−1.321)		(−0.480)		(−0.777)		(−0.194)
dollarGlobal		0.385^{**}		0.576^{**}		0.787^{*}		1.170
		(2.468)		(2.345)		(1.880)		(1.482)
FG		0.000		0.006		0.015		0.030
		(0.026)		(0.973)		(1.235)		(1.365)
HPW		−0.001		−0.004		−0.020		-0.051^{**}
		(−0.277)		(−0.407)		(−1.398)		(−2.116)
WZ		0.001		−0.006		0.002		0.013
		(0.242)		(−0.567)		(0.111)		(0.442)
VIX		−0.006		−0.012		−0.028		−0.029
		(−1.022)		(−1.086)		(−1.382)		(−0.904)
VRP		0.006		0.009		0.017		0.019
		(1.401)		(1.225)		(1.483)		(0.968)
MOVE		−0.004		−0.004		−0.009		−0.017
		(−1.290)		(−0.897)		(−0.961)		(−1.117)
R2 adj	0.062	0.153	0.038	0.105	0.091	0.194	0.158	0.292

Table 8: Excess return regression (22) applied to Eurodollar deposits. The left-hand side is the monthly excess return of holding a 2-, 3-, 6-, or 12-month Eurodollar deposit for 1 month. The PCs are normalized to have mean zero. Variables FG, HPW, WZ, VIX, VRP, and MOVE are normalized to have mean zero and standard deviation one. All the regression coefficients are multiplied by 100. The t-statistics are reported in parentheses, and statistical significance at the 10%, 5%, and 1% levels are denoted by *, **, and ***, respectively.

	2-month Eurodollar		3-month Eurodollar		6-month Eurodollar		12-month Eurodollar	
Const	0.006^{***}	-0.029	0.011^{***}	-0.029	0.036^{***}	-0.132	0.044^{***}	-0.423
	(5.298)	(−1.378)	(5.212)	(−0.783)	(6.378)	(−1.289)	(2.843)	(−1.623)
goldPC1		0.067		0.099		0.277		1.240
		(0.487)		(0.423)		(0.448)		(0.794)
goldPC2		0.038		0.031		0.743		1.126
		(0.090)		(0.041)		(0.326)		(0.190)
EDPC1	0.020	−0.016	0.076^{*}	0.088	0.177	0.333	0.384	1.411^{**}
	(1.022)	(−0.302)	(1.891)	(0.867)	(1.452)	(1.221)	(1.013)	(2.099)
EDPC2	0.131	−0.448	−0.262	−0.496	−0.558	−0.620	0.365	4.438
	(0.277)	(−0.794)	(−0.277)	(−0.458)	(−0.217)	(−0.217)	(0.056)	(0.611)
Einf		0.735^{*}		0.559		3.673		7.968
		(1.839)		(0.719)		(1.637)		(1.309)
Ergdp		0.130		0.099		0.702		2.782
		(0.284)		(0.128)		(0.330)		(0.512)
goldRet		0.006		0.009		0.042		−0.189
		(0.266)		(0.198)		(0.303)		(−0.503)
goldVol		−0.017		−0.049		−0.130		−0.313
		(−0.848)		(−1.335)		(−1.299)		(−1.277)
DivYield		5.164^{*}		10.755^{*}		13.996		31.831
		(1.686)		(1.960)		(1.067)		(0.983)
EquityVW		−0.035		−0.061		−0.167		−0.486
		(−0.982)		(−0.920)		(−0.944)		(−1.015)
SMB		−0.019		0.023		−0.019		−0.076
		(−0.604)		(0.431)		(−0.131)		(−0.198)
HML		−0.058		−0.026		0.033		0.332
		(−1.091)		(−0.273)		(0.146)		(0.580)
Baa-Aaa		0.720		1.154		4.797		18.713
		(1.119)		(0.899)		(1.287)		(1.582)
FXRet		−0.094		−0.123		0.242		1.724
		(−0.758)		(−0.491)		(0.353)		(0.929)
FXHML		−0.042		0.074		0.085		0.096
		(−0.764)		(0.711)		(0.340)		(0.155)
dollarGlobal		−0.043		-0.197^{*}		−0.318		−0.011
		(−0.723)		(−1.662)		(−0.902)		(−0.011)
FG		0.003		0.003		0.002		−0.018
		(1.623)		(0.821)		(0.184)		(−0.748)
HPW		0.004^{**}		0.007^{**}		0.012		0.007
		(2.260)		(2.421)		(1.356)		(0.325)
WZ		-0.006^{***}		-0.011^{***}		-0.033^{***}		-0.084^{***}
		(−3.884)		(−3.781)		(−4.099)		(−4.076)
VIX		−0.003		−0.004		−0.008		−0.028
		(−1.482)		(−0.882)		(−0.708)		(−0.853)
VRP		0.001		0.000		0.005		0.034
		(0.735)		(0.008)		(0.696)		(1.608)
MOVE		−0.002		-0.004^{*}		-0.012^{*}		−0.023
		(−1.442)		(−1.690)		(−1.842)		(−1.421)
R2 adj	−0.007	0.172	0.000	0.160	−0.003	0.199	−0.006	0.195

6 Conclusion

In this paper we investigate the dynamics of gold lease rates, interests paid in gold for borrowing gold. We find that risks in both the level and the slope of gold interest rates are strongly priced. The risk premia are increasing in the slope of the gold interest-rate curve. Over our sample period, the slope is highly time-varying, mostly positive but occasionally negative. The time-varying slope and associated risk premia lead to a strong rejection of the expectation hypothesis. In contrast, the expectation hypothesis cannot be consistently rejected for comparable US interest rate data over the same sample period. We further show that excess returns in gold bonds are also significantly predicted by gold volatility and returns in equity, credit and currency markets, in addition to the first two PCs of the gold lease rates.

As central banks around the world continue to hold gold as part of their reserve assets, and the financial markets increasingly accept gold as eligible collateral, it seems all the more important to better understand the risk-return dynamics in the gold lending market. While our analysis provides initial evidence of risk premium behavior in this market, many questions remain open. For example, what are the economic contents of the slope factor that make it so important in determining the risk premia in gold lending market? Moreover, can risk premia in gold lease rates help predict returns of other asset classes and macroeconomic conditions? Finally, since gold is not tied to any specific currency or sovereign country, to what extent do gold interest rates reflect the (shadow) real interest rates for the global economy? We leave these questions for future research.

A Robustness Check Using Futures Data

So far, we have tested the expectations hypothesis and characterized risk premium in gold lease rates implied by GOFO. In this section, we reproduce our key results using futures data. We find the results to be robust to whether GOFO or futures are used to construct the gold lease rates.

A.1 The Futures Data

The futures data we use are from the CME Group. Each CME gold futures contract is equivalent to 100 ounces of gold. CME specifies the trading and delivery of gold futures as as follows: "Trading is conducted for delivery during the current calendar month; the next two calendar months; any February, April, August, and October falling within a 23-month period; and any June and December falling within a 72-month period beginning with the current month. ... Trading terminates on the third last business day of the delivery month. ... Delivery may take place on any business day beginning on the first business day of the delivery month or any subsequent business day of the delivery month, but not later than the last business day of the current delivery month."[26]

The key implications of the futures contract specification are: (i) futures data are not constant maturity, and, more importantly, (ii) each futures contract can be delivered as early as the start of a month and as late as the end of the month.[27] The latter feature introduces substantial uncertainty regarding the effective maturity of a gold futures contract and, together with it, substantial measurement errors in futures-implied gold lease rates. This problem is particularly acute at short maturities.

To address the uncertainty associated with futures data, we use the following method and convention. We take the first week of each delivery month as the maturity date of the corresponding futures contract. We then exclude all contracts that have maturity less than 6 weeks because these contracts are most sensitive to the assumed maturity date. We fill in the short end of the futures curve by including the 4-week GOFO. Futures contract that mature longer than 12 months are also excluded, in order to be consistent with GOFO data. In addition, because the futures prices is proportional to the spot gold price, and because spot gold prices are too noisy to measure, we use the forward rates implied by the futures

[26] See http://www.cmegroup.com/trading/metals/precious/gold_contract_specifications.html.

[27] See http://www.cmegroup.com/delivery_reports/MetalsIssuesAndStopsMTDReport.pdf for monthly delivery report of CME gold futures. This report shows the dispersion of delivery dates within the current delivery month.

term structure as the input to our model. Finally, we will ignore the daily mark-to-market of the futures contracts; this means that the futures prices and forward prices should be equal, if it were not for measurement errors and the delivery option in the futures market. For simplicity, in the exposition below we call our final data set consisting of gold futures and 4-week GOFO the "futures data."[28]

A.2 Comparison between Futures Data and GOFO Data

Having described the procedure of cleaning the futures data, we now show a few comparisons of gold lease rates implied by the two data sources: futures data and GOFO data. Table 9 presents the correlations between the GOFO-implied gold lease rates and the futures-implied gold lease rates. The left half shows the spot-rate correlations, and the right half shows the forward-rate correlations. These correlations are high (at least 0.98 for spot rates and at least 0.95 for forward rates). Since the inclusion of 4-week GOFO data cannot affect the forward rates beyond 4 weeks, these high correlations are not mechanical.

Table 9: Correlations between gold lease rates implied by GOFO and filtered from futures data. The left half shows the correlations of the spot rates, and the right half shows the correlations of the forward rates.

Spot rate correlations					
	4w	9w	13w	26w	52w
4w	0.979	0.971	0.962	0.925	0.848
9w	0.994	0.991	0.986	0.958	0.893
13w	0.993	0.995	0.993	0.974	0.920
26w	0.960	0.975	0.983	0.989	0.968
52w	0.891	0.920	0.938	0.971	0.986

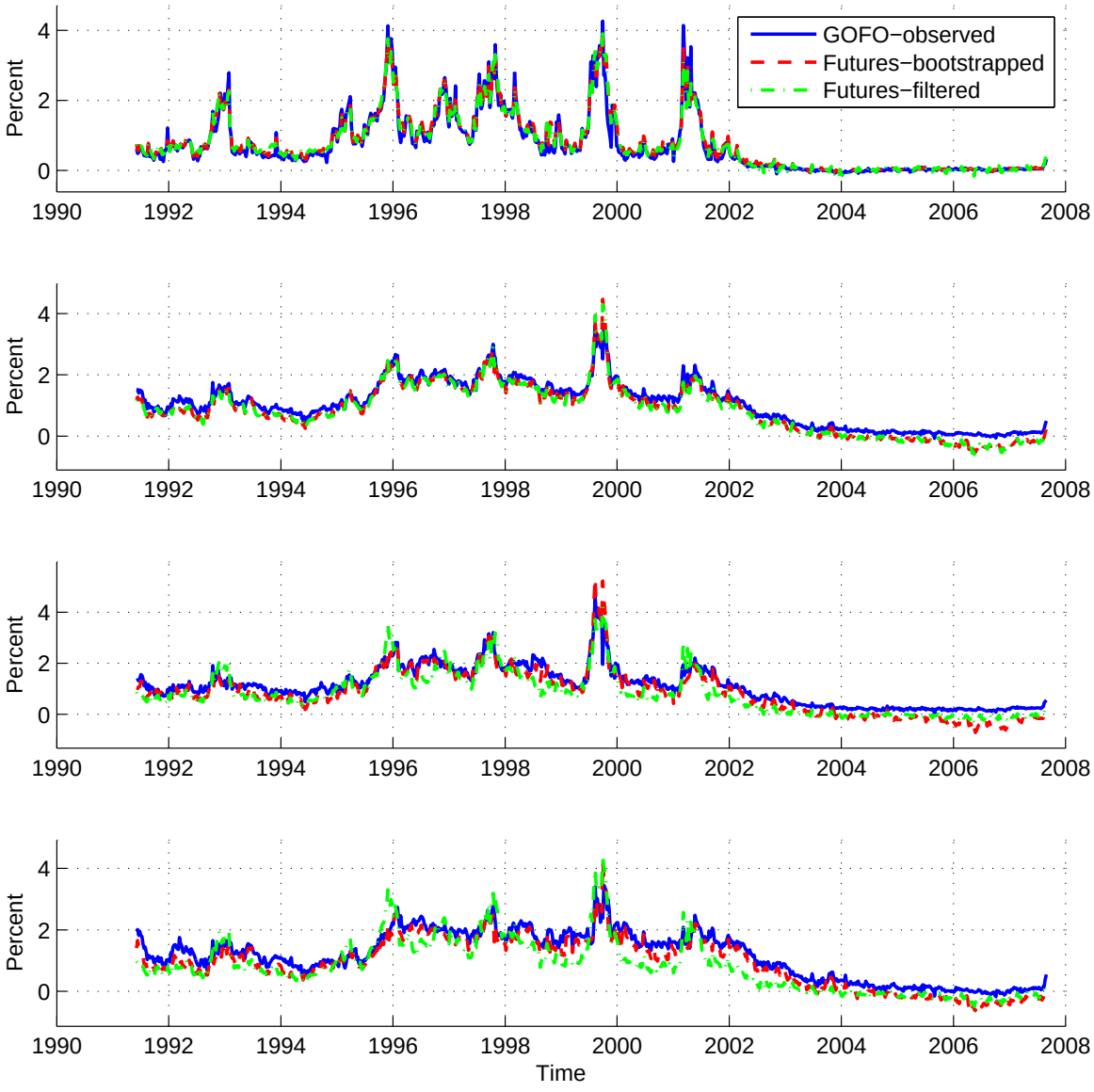
Forward rate correlations					
	4w	4-9w	9-13w	13-26w	26-52w
4w	0.979	0.961	0.933	0.874	0.740
4-9w	0.985	0.984	0.970	0.930	0.823
9-13w	0.930	0.949	0.951	0.935	0.865
13-26w	0.877	0.924	0.950	0.967	0.947
26-52w	0.774	0.844	0.890	0.937	0.963

Figure 8 shows the time-series of the gold lease rates (spot and forward) computed using three different methods: by directly computing from GOFO data (blue solid lines), by bootstrapping the futures data using constant forward rates (red dashed lines), and by filtering the futures data. We observe that these three time series closely track each other. A

[28] We illustrate our method of data cleaning by an example. Suppose that in the first week of August 2013, we observe the raw prices of futures contracts that mature on the third last business day of August 2013, September 2013, October 2013, November 2013, December 2013, February 2014, April 2014, June 2014, and August 2014. These futures mature in 0, 1, 2, 3, 4, 6, 8, 10, and 12 months. First, we exclude the first two contracts, August 2013 and September 2013, because they mature within 6 weeks. Second, we include the 4-week GOFO data. Third, we calculate the 2-3, 3-4, 4-6, 6-8, 8-10, and 10-12 months forward rates. Our final input to the model includes those forward rates as well as the 4-week zero rate from GOFO.

slightly subtler pattern is that futures-implied lease rates tend to be lower than GOFO-implied lease rates, especially for longer maturities. This pattern can come from the delivery option of the short side in the futures market.

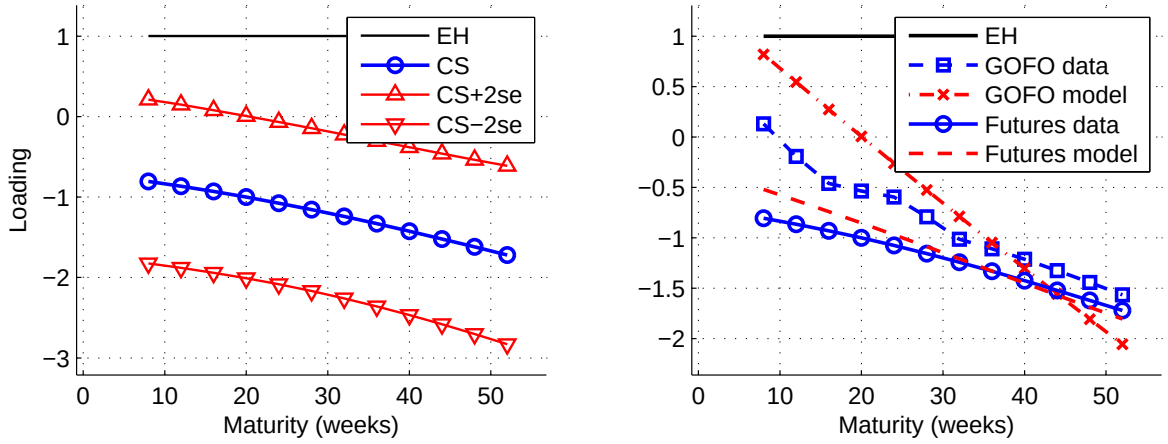
Figure 8: Comparison between filtered and observed estimates of gold lease rates, spot and forward. First plot: 4-week spot rates. Second plot: 52-week spot rates. Third plot: 13-26 week forward rates. Fourth plot: 26-52 week forward rates.



A.3 Campbell-Shiller Regression

Figure 9 reports the results of Campbell-Shiller regression specified in Section 3. As we can see from the left panel, the coefficients are significantly below unity and decrease in maturity, indicating a rejection of the expectations hypothesis. The right panel shows that the Campbell-Shiller regression coefficients implied by data and model, for both the futures data and GOFO data, are close to each other and all below unity. The coefficients for futures data are even lower than those for the GOFO data. That is, the EH is rejected by the Campbell-Shiller test in the futures data.

Figure 9: Campbell-Shiller regression coefficients



(a) Futures data

(b) Comparing futures and GOFO data

A.4 Term Structure Estimation

Table 10 presents the parameter estimations of a two-factor term structure model using futures-implied gold lease rates. Again, the parameter estimates are close to their counterparts using GOFO-implied gold lease rates (see Table 2 and Table 3 of Section 4). The market price of risk parameters are significantly negative on the slope, implying that a higher slope of the gold lease rates is associated with a higher excess return. Again, this predictability of excess returns rejects the EH.

Table 10: Parameter estimates of a two-factor term structure model using futures data. The top panel shows the feedback matrices and covariance matrix, and the bottom panel shows the market price of risk parameters.

	$K_1^{\mathbb{P}}$		$K_1^{\mathbb{Q}}$		Σ	
Ests	0.977^{***}	-0.006	1.000^{***}	0.152^{***}	0.419^{***}	
	0.009^{***}	0.946^{***}	-0.005^{***}	0.972^{***}	-0.086^{***}	0.115^{***}
s.e.	0.008	0.046	0.001	0.001	0.012	
	0.003	0.009	0.001	0.001	0.003	0.004

	VAR risk		χ^2	STD risk		χ^2
Ests	0.013	-1.803^{***}	0.000	-0.054^{***}	-0.376^{***}	0.000
	0.692^{**}	-4.407^{***}		0.080^{**}	-0.508^{***}	
s.e.	0.106	0.589		0.018	0.115	
	0.345	1.334		0.038	0.155	

B Estimation of No-Arbitrage Term Structure Models

In this section, we show how to analytically concentrate out $\kappa_0^{\mathbb{P}}$, $\kappa_1^{\mathbb{P}}$, σ_e, and $r_\infty^{\mathbb{Q}}$ for the WOE estimation.

B.1 Notations

Continuing from Section 4, we write:

$$A_{n,X} = r_\infty^{\mathbb{Q}} + tr(\Sigma_X C_{n,X}), \quad \text{where } C_{n,X} = -\frac{1}{2n}\sum_{i=0}^{n-1} i^2 B_{i,X} B_{i,X}'. \tag{23}$$

It is important to note that both $B_{n,X}$ and $C_{n,X}$ are only functions of $\lambda^{\mathbb{Q}}$.

Obviously, $B_{n,\mathcal{P}}$ is only dependent on $\lambda^{\mathbb{Q}}$. To see clearly how $r_\infty^{\mathbb{Q}}$, $\Sigma_{\mathcal{P}}$ and λ^{Q} influence

$A_{n,\mathcal{P}}$, we first write:

$$\begin{aligned} A_{n,X} =& r_\infty^{\mathbb{Q}} + tr(\Sigma_X C_{n,X}) && (24)\\ =& r_\infty^{\mathbb{Q}} + tr(V^{-1}\Sigma_{\mathcal{P}}(V')^{-1}C_{n,X}) && (25)\\ =& r_\infty^{\mathbb{Q}} + tr(\Sigma_{\mathcal{P}}(V')^{-1}C_{n,X}V^{-1}) && (26)\\ =& r_\infty^{\mathbb{Q}} + vech((V')^{-1}C_{n,X}V^{-1})vec(\Sigma_{\mathcal{P}}) && (27)\\ =& r_\infty^{\mathbb{Q}} + D_{n,X}vec(\Sigma_{\mathcal{P}}), && (28) \end{aligned}$$

where $D_{n,X} = vech((V')^{-1}C_{n,X}V^{-1})$ and $V = WB_X$.

Stacking $A_{n,X}$ for all maturities used in the estimation, we obtain:

$$A_X = \iota_J r_\infty^{\mathbb{Q}} + D_X vec(\Sigma_{\mathcal{P}}), \tag{29}$$

where D_X is obtained by staking the relevant row vectors $D_{n,X}$ on top of one another. ι_J denotes a J-size vector of ones.

Now plugging $A_{n,X}$ and A_X into $A_{n,\mathcal{P}}$, we have:

$$\begin{aligned} A_{n,\mathcal{P}} =& r_\infty^{\mathbb{Q}} + D_{n,X}vec(\Sigma_{\mathcal{P}}) - B'_{n,X}V^{-1}W(\iota_J r_\infty^{\mathbb{Q}} + D_X vec(\Sigma_{\mathcal{P}})) && (30)\\ =& (1 - B'_X V^{-1}W\iota_J)r_\infty^{\mathbb{Q}} + (D_X - B'_X V^{-1}WD_X)vec(\Sigma_{\mathcal{P}}) && (31)\\ =& E_n r_\infty^{\mathbb{Q}} + tr(\Sigma_{\mathcal{P}}F_n), && (32) \end{aligned}$$

where

$$E_n = 1 - B'_{n,X}V^{-1}W\iota_J, \tag{33}$$

and F_n is obtained by appropriately collapsing the row vector $D_{n,X} - B'_{n,X}V^{-1}WD_X$ into a square matrix. It is important to note that both E_n and F_n are only dependent on $\lambda^{\mathbb{Q}}$.

Stacking $A_{n\mathcal{P}}$ for all maturities n used in estimation, we obtain:

$$A_{\mathcal{P}} = E r_\infty^{\mathbb{Q}} + F, \tag{34}$$

where E is obtained by stacking up E_n's, and F is obtained by stacking up $tr(\Sigma_{\mathcal{P}}F_n)$, for relevant n's.

B.2 $\kappa_0^{\mathbb{P}}$ and $\kappa_1^{\mathbb{P}}$

We know from JSZ that the globally optimal estimates of $\kappa_0^{\mathbb{P}}$ and $\kappa_1^{\mathbb{P}}$ can be obtained from a VAR of $\mathcal{P}_{t+1}$ on $\mathcal{P}_t$:

$$\kappa_1^{\mathbb{P}} = E_T[\mathcal{P}_{t+1}^e(\mathcal{P}_t^e)'](E_T[\mathcal{P}_t^e(\mathcal{P}_t^e)'])^{-1}, \tag{35}$$

$$\kappa_0^{\mathbb{P}} = E_T[\mathcal{P}_t] - \kappa_1^{\mathbb{P}} E_T[\mathcal{P}_t], \tag{36}$$

where $E_T[.]$ denotes sample average and $\mathcal{P}_t^e = \mathcal{P}_t - E_T[\mathcal{P}_t]$.

B.3 σ_e

Likewise, the global estimate of σ_e can be obtained from the first order condition of the likelihood function:

$$\sigma_e^2 = \frac{1}{T(J-N)} \sum_t ||\mathcal{P}_{t+1}^e - W^e(A_{\mathcal{P}} + B_{\mathcal{P}}\mathcal{P}_{t+1})||_2. \tag{37}$$

B.4 $r_\infty^{\mathbb{Q}}$

The first order condition with respect to $r_\infty^{\mathbb{Q}}$ is:

$$\left(E_T[\mathcal{P}_{t+1}^e] - W^e(A_{\mathcal{P}} + B_{\mathcal{P}}E_T[\mathcal{P}_{t+1}])\right)' W^e \frac{\partial A_{\mathcal{P}}}{\partial r_\infty^{\mathbb{Q}}} = 0. \tag{38}$$

Utilizing the earlier representation, $A_{\mathcal{P}} = Er_\infty^{\mathbb{Q}} + F$, we have

$$\left(E_T[\mathcal{P}_{t+1}^e] - W^e(F + B_{\mathcal{P}}E_T[\mathcal{P}_{t+1}]) - W^e E r_\infty^{\mathbb{Q}}\right)' W^e E = 0, \tag{39}$$

which is linear in $r_\infty^{\mathbb{Q}}$. Collecting terms, we have:

$$r_\infty^{\mathbb{Q}} = \frac{\left(E_T[\mathcal{P}_{t+1}^e] - W^e(F + B_{\mathcal{P}}E_T[\mathcal{P}_{t+1}])\right)' W^e E}{||W^e E||_2}. \tag{40}$$

References

Andrews, D., Monahan, C., 1992. An improved hac covariance matrix estimator. Econometrica 60, 953–966.

Ang, A., Bekaert, G., Wei, M., 2007. Do macro variables, asset markets or surveys forecast inflation better?. Journal of Monetary Economics 54(4), 1163–1212.

Backus, D. K., Zin, S. E., 1993. Long-memory inflation uncertainty: evidence from the term structure of interest rates. Journal of Money, Credit and Banking 25, 681–700.

Bailey, W., Chan, K., 1993. Macroeconomic influences and the variability of the commodity futures basis. Journal of Finance 48(2), 555–573.

Barro, R. J., Misra, S. P., 2013. Gold returns. Working paper NBER.

Basel II, 2006. Basel II: International convergence of capital measurement and capital standards: A revised framework – comprehensive version. Basel Committee on Banking Supervision, Bank for International Settlements.

Basel III, 2012. Basel III counterparty credit risk and exposures to central counterparties – frequently asked questions. Basel Committee on Banking Supervision, Bank for International Settlements.

Bollerslev, T., Tauchen, G., Zhou, H., 2009. Expected stock returns and variance risk premia. Review of Financial Studies 22, 4463–4492.

Buraschi, A., Menini, D., 2002. Liquidity risk and specialness. Journal of Financial Economics 64, 243–284.

Campbell, J., Shiller, R., 1991. Yield spreads and interest rate movements: A bird's eye view. Review of Economic Studies 58, 495–514.

Casassus, J., Collin-Dufresne, P., 2005. Stochastic convenience yield implied from commodity futures and interest rates. Journal of Finance 60, 2283–2331.

Chen, R., Scott, L., 1993. Maximum likelihood estimation for a multifactor equilibrium model of the term structure of interest rates. Journal of Fixed Income 3, 14–31.

Cochrane, J., Piazzesi, M., 2005. Bond risk premia. American Economic Review 95, 138–160.

Dai, Q., Singleton, K., 2002. Expectations puzzles, time-varying risk premia, and affine models of the term structure. Journal of Financial Economics 63, 415–441.

Downing, C., Oliner, S., 2007. The term structure of commercial paper rates. Journal of Financial Economics 83, 59–86.

Erb, C. B., Harvey, C. R., 2013. The golden dilemma. Financial Analysts Journal 69(4), 10–42.

Fama, E., French, K., 1987. Commodity futures prices: Some evidence on forecast power, premiums, and the theory of storage. Journal of Business 60, 55–73.

——— , 1992. The cross-section of expected stock returns. Journal of Finance 47, 427–465.

Fama, E. F., Bliss, R. R., 1987. The information in long-maturity forward rates. American Economic Review 77(4), 680–692.

Fontaine, J.-S., Garcia, R., 2012. Bond liquidity premia. Review of Financial Studies 25, 1207–1254.

Hansen, L., 1982. Large sample properties of generalized method of moments estimators. Econometrica 50, 1029–1054.

Hu, X., Pan, J., Wang, J., 2012. Noise and information for illiquidity. Working paper MIT.

Joslin, S., Le, A., 2011. Interest rate volatility and no-arbitrage term structure models. Unpublished working paper. UNC and USC.

Joslin, S., Le, A., Singleton, K., 2012. Why gaussian macro-finance term structure models are (nearly) unconstrained factor-vars. Journal of Financial Economics, forthcoming.

Joslin, S., Singleton, K., Zhu, H., 2011. A new perspective on gaussian dynamic term structure models. Review of Financial Studies 24, 926–970.

Jotikasthira, P., Le, A., Lundblad, C., 2013. Why do term structure in different currencies comove?. Unpublished working paper. UNC.

Lange, J., Sack, B., Whitesell, W., 2003. Anticipations of monetary policy in financial markets. Journal of Money, Credit, and Banking 35, 889–909.

Le, A., Singleton, K., Dai, J., 2010. Discrete-time affine^q term structure models with generalized market prices of risk. Review of Financial Studies 23, 2184–2227.

Longstaff, F., 2000. The term structure of very short-term rates: New evidence for the expectation hypothesis. Journal of Financial Economics 58, 397–415.

Lustig, H., Roussanov, N., Verdelhan, A., 2011. Common risk factors in currency markets. Review of Financial Studies 24(11), 3731–3777.

Neuberger, A., 2001. Gold derivatives: The market impact. Report prepared for the world gold council London Business School.

Newey, W., West, K. D., 1987. A simple positive semi-definite, heteroskedasticity and autocorrelation consistent covariance matrix. Econometrica 55, 703–708.

Piazzesi, M., Swanson, E., 2008. Futures prices as risk adjusted forecasts of monetary policy. Journal of Monetary Economics 55, 677–691.

Schwartz, E. S., 1997. The stochastic behavior of commodity prices: Implications for valuation and hedging. Journal of Finance 52(3), 923–973.

Verdelhan, A., 2013. The share of systematic variation in bilateral exchange rates. Working paper MIT.

World Gold Council, 2013a. Central bank diversification strategies: Rebalancing from the dollar and euro. .

——— , 2013b. World official gold holdings. .

Wright, J., Zhou, H., 2009. Bond risk premia and realized jump risk. Journal of Banking and Finance 33, 2333–2345.