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Information Risk in TIPS Market: An Analysis of Nominal and Real Interest Rates

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Abstract. This study investigates the presence of information risk in two closely linked interest rate securities traded in separate markets: the nominal interest rate observed in the Treasury bond market and the real interest rate observed in the relatively new Treasury Inflation-Protected Securities (TIPS) market. We find that information flows unilaterally from the Treasury bond market to the TIPS market with a one-day lag. The information risk arising from asymmetric information flows may cause less informed traders to demand a higher rate of return (O'Hara, 2003). Our study provides an empirical explanation of why the TIPS yield has been relatively high throughout its nascent trading history.

Key words: Treasury inflation-protected securities, real interest rate, information risk, vector error correction model

JEL Classification: G14, E43, C32

1. Introduction

Treasury Inflation-Protected Securities (TIPS) are becoming a more important component of Treasury debt. The surprisingly high yields of TIPS, relative to conventional Treasuries, have been both a puzzle and a concern to the U.S. Treasury Department due to the excess interest cost (Sack and Elsasser, 2004). When initiated in 1997, TIPS were expected to lower the borrowing cost of the Treasury due to the elimination of an inflation risk premium. Over the last seven years the Treasury has paid an estimated $3 billion more in interest by issuing TIPS instead of nominal Treasuries with a comparable maturity (Sack and Elsasser, 2004). The difference between the nominal yield of nominal bonds and the real yield of comparable maturity TIPS was also expected to inform monetary policymakers of the market's inflation expectations, yet the difference has been consistently below actual inflation throughout its trading history.

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Sack (2000), Shen and Corning (2001), and Sack and Elsasser (2004) focus on the possible reasons for a “too high” TIPS real yield and consider a liquidity premium, a lower supply, and a lack of knowledge about the new TIPS as possible reasons. However, as Sack and Elsasser (2004) conclude, liquidity has improved over the past seven years, the supply has increased, and there is wider acceptance of the securities, yet the relatively high TIPS real yield remains. Anecdotal evidence remains that traders in TIPS may be at a disadvantage versus the nominal Treasury market traders due to a turnover rate that is roughly one half of off-the-run nominal Treasuries. In addition, the evidence of bid-ask spreads reveals an average of two ticks for TIPS maturities of five to ten years, compared to one-half to one tick for nominal Treasuries. However, neither of these proxies for liquidity differences is large enough to explain a 50 basis point difference between the ten-year inflation expectations reported by the Survey of Professional Forecasters, conducted by the Federal Reserve Bank of Philadelphia, and the TIPS-nominal bond spread from 1997–2003 (Sack and Elsasser, 2004). This same study found that after eliminating the 1997–1999 period which included bid ask spreads exceeding those of off-the-run securities, the unexplained difference remained at 45 basis points. A one tick difference in the spread is not close to explaining a 45 basis point difference. Until this valuation puzzle is explained, TIPS will not be entirely useful in isolating the market’s implied inflation expectations.

While much of the focus on the puzzle to date has concerned liquidity, an alternative explanation of the higher TIPS yield is information risk, or the risk of price discovery. O’Hara (2003) suggests that information risk be incorporated into asset valuation. According to O’Hara, the more trading that occurs from traders with private information, called informed traders, the more information risk is perceived by uninformed traders. Her asymmetric information asset pricing model demonstrates that if some traders find themselves disadvantaged in obtaining the most current information, they will demand higher compensation and set a lower price for bearing this information risk.

While her 2003 study focuses on stocks with public and private information, and a recent paper (Easley and O’Hara, 2004) investigates the impact of information risk on a firm’s cost of capital, information risk can also exist in interest rate markets. Just as one stock may have both informed and uninformed traders, the real interest rate may have both types of traders. As new information impacting the real interest rate arrives continuously, primary dealers, with access to more information than secondary dealers, fit the description of informed traders with private information as described by Umlauf (1991). It is also likely that traders better informed about economic news are attracted to the liquidity of the enormous nominal Treasury bond market as well as to the highly developed interest rate derivative markets as they act on superior information.

The Grossman-Stilglitz (1980) model indicates that uninformed traders know there are informed traders, but not the information that informed traders have; therefore, they make inferences from the price itself and the fully-revealing equilibrium price emerges. In explaining the difference in her asymmetric pricing model relative to the Grossman-Stiglitz model, O’Hara (2003) says, “The innovation here is the argument that when information is asymmetric, uninformed investors demand compensation for portfolio-induced risks which they cannot diversify.”

Both nominal Treasury and TIPS markets share a common component—the expected real return—but they contain other return components that behave very differently in the presence of inflation, and therefore attract different clienteles who must hold non-replicating portfolios. Investors holding TIPS are attracted to its inflation hedging properties. However, they cannot diversify into other securities without giving up the inflation hedging properties. If there is information risk by holding the TIPS, because the real rate information is likely to be revealed first in the nominal Treasury market, they will need additional compensation for taking this risk. Moreover, changes in real rates of interest create a more volatile price for TIPS than for nominal Treasuries due to longer real durations for TIPS (Roll, 2004), so perception of an informational disadvantage is a consequential risk.

Our study explores whether TIPS traders may fit the description of less informed traders, who are aware that new information about the real rate of interest is first likely to occur in the nominal Treasury bond market. Sack and Elsasser (2004) find that primary dealers are far less active in the TIPS market than in the nominal Treasury market where they dominate the trading. If the participants of the TIPS market perceive they are at an informational disadvantage relative to participants in the nominal Treasury market, then the TIPS investors will demand additional compensation for the same reason as the less informed stock investors in O'Hara (2003).

Easley and O'Hara (2004) explain the difficulty of measuring the dominance of informed versus uninformed traders in a particular stock. The presence of two interest rate markets, both pricing the real interest rate, creates an opportunity to identify the less informed market as measured by the direction of information flow as well as the lag in price innovation. Once identified, according to O'Hara's asymmetric information asset pricing model, the disadvantaged market will demand a higher compensation. In fact, we find that TIPS investors are at a disadvantage in price discovery with a one-day lag. While it is possible that informed traders could transact their trades in either market, the evidence is that information flows unilaterally from the Treasury bond market to the TIPS market.

The reason for such a lag in price innovation is also analogous to Easley and O'Hara's (2004) description of stocks with more public information having a "greater institutional following." Bond market participants who have private information about the real rate of interest could act on this information either in the nominal Treasury market or in the TIPS market. Our findings are that they choose to act in the former. This is likely because the nominal Treasury market, with its highly developed infrastructure and large number of traders in both the spot and derivative markets, is the most profitable vehicle for trading on new information about expected real interest rates and thus aggregating it into the price, making information public. However, it is beyond the scope of our study to offer evidence of why informed participants choose the nominal market; we only document that they do choose it.

In a perfectly integrated system, both the nominal and the real rate markets, for securities with the same maturity, impound information instantaneously, so that prices adjust to a new equilibrium with no lag. According to revised Fisher models under uncertain inflation, such as in Benninga and Protopapadakis (1983), as new information arrives, both markets adjust so that the spread between the nominal yield offered and the real yield offered provides compensation for both expected inflation and an inflation risk premium. However, as O'Hara

(2003) points out, often the facts do not confirm a perfectly integrated system. Different market structures, security designs, and trading activities uniquely affect the processing and absorption of incoming information. Consequently, one market may lead the other in the process of incorporating information, even though the two rates, and the spread between them, are driven by common economic factors.

The objective of our study is to present empirical evidence about the presence of information risk in TIPS market. We use daily price histories of TIPS and Treasury STRIPS from Datastream and find that the ten-year spot real and nominal rates are cointegrated. We then apply a Granger causality test to measure the short term lead-lag relationship. Next we use Gonzalo and Granger's (1995) common component analysis to test for the dominant contributor to the permanent component of price innovations between these two cointegrated interest rates. Finally, we apply vector error correction models (VECM) to identify the price adjustment process in the two markets.

Our study documents that the error correction process occurs primarily in the TIPS market. It takes one additional day to complete the adjustment process for the deviation from the long term relationship between nominal and real rates. We also document a unidirectional flow of information from the nominal Treasury market to the TIPS market. This finding supports the presence of information risk in the TIPS market and provides support for O'Hara's assertion that asymmetric information requires compensation when the market realizes it is relatively uninformed. The information risk becomes an explanation for a higher TIPS yield.

2. Literature review and hypotheses

After seven years of trading history, empirical studies on the pricing behavior of TIPS have yielded some facts about this relatively new inflation hedge security. The earlier research focused on the TIPS yield as a measure of the ex ante real rate for the purpose of extracting the unobservable expected inflation rate from the conventional Treasury yield. Sack (2000) and Shen and Corning (2001) use TIPS to derive inflation expectations by subtracting the yield on TIPS from a STRIP comparable to maturity after adjusting for the "off the run" spread assumed to be associated with the TIPS. Sack (2000) finds that the variation in the ten-year inflation outlook is higher than would be expected, and Shen and Corning (2001) attribute the underestimation of inflation to a liquidity premium imbedded in the TIPS yield.

Sack and Elsasser (2004) study the bid-ask spreads of TIPS and off-the-run Treasury securities and conclude that "TIPS liquidity has improved much in recent years and is currently not far below that of off-the-run nominal Treasuries." They raise the possibility that a low relative valuation of the TIPS is a result of a decline in near-term inflationary pressures during the last three years. Chu, Pittman and Yu (2004) investigate the prices of the first maturing TIPS issue during its last coupon period, and find that during the first quarter of 2002 investors required a premium to hold TIPS instead of a Treasury bill with the same maturity.

A second subset of empirical research has involved the Fisherian relationship between ex ante real yield and the ex ante nominal yield. The price history of TIPS enables these tests to be conducted with far fewer assumptions than were previously required in the absence of an inflation-hedge security. Chu, Pittman and Yu (2003) extract the ex ante real pure discount rate with a constant ten-year maturity from TIPS prices. They find that the time series of real and nominal spot rates are cointegrated. The cointegrated system casts doubt on the accuracy of tests of the Fisher effect that infer a constant or stationary real rate. Laatsch and Klein (2003) also find that both the nominal and TIPS yields are nonstationary at order one, and that the two series are cointegrated.

A third area of empirical inquiry is Roll's (2004) study of the TIPS as a new asset class in a diversified portfolio. Roll (2004) finds that daily returns of long-term nominal bonds are strongly and positively correlated with comparable-maturity TIPS, and both are negatively correlated to equity returns. He also finds that TIPS nominal return volatility is less than the volatility of conventional bonds, and that TIPS return volatility is time varying. Roll's study explains why there is considerable interest in the behavior of TIPS prices.

An empirical study that has not yet been undertaken for the relatively new inflation hedge is the impact of information risk on the prices of TIPS. Exploring the way that the nominal and TIPS prices evolve in markets can shed light on whether there is a premium for information risk imbedded in the TIPS real yield. O'Hara (2003) asserts that the nature of the information arrival process is important in determining an asset's price. Instead of using off-the-run spreads or trading volume as a measure of liquidity, it is more informative to study precisely how security prices adjust in response to the arrival of information. If new information is always derived from one market to another, then the traders in the second market will demand compensation for information risk. If information is asymmetric with new information originating in the nominal bond market and then flowing to the TIPS market, where there are fewer informed traders, Easley and O'Hara (2004) would expect the asset yield to be higher in the TIPS market.

The vector error correction model (VECM) has been applied to study information flow among informationally linked markets. Harris, McInish and Wood (2002) use the methodology on Dow Jones 30 stocks traded on three national stock exchanges. They identify multilateral information flows among the three stock exchanges. Chatrath, Chaudhry and Christie-David (1999) investigate the information flow between Eurodollar futures and Treasury bill futures. They find that information flows in both directions during the error correction process. The findings of bilateral or multilateral information flows are consistent with informationally linked transparent equity and futures markets. Unlike the centralized price reporting system adopted by equity and futures markets, trading of Treasury securities occurs on various brokerage trading desks. This study sheds light on the information flow between two informationally linked but less transparent markets.

We expect to find that information flows from the nominal Treasury bond market, although we do not have an a priori view of the length of time before new information is incorporated into the TIPS prices. Sack and Elsasser (2004) describe an inability of investors to adjust to this new asset, a lack of major dealer participation, low liquidity, and differences in supply, which indicate that the lag in the transmission of information may be quite significant. Transmission of new information concerning the real rate can be readily transmitted throughout the nominal Treasury markets via a vast derivatives market in nominal interest rates; however, we expect the TIPS market would be more reactionary, observing first the changes in the nominal yields before impounding the portion of the information that impacts the real yield.

Table 1. Summary statistics (Time period: April 8, 1999–September 7, 2001)

	Coupon STRIPS	TIPS
Panel A: Sample data statistics		
Average number of securities per day	93	7
Minimum number of securities per day	89	6
Maximum number of securities per day	95	8
Total number of security prices	58,768	4,390
Total number of observation days	632	632

	Basic statistics (in percentage)				Autocorrelation					
	Mean	S.D.	Maximum	Minimum	ρ_1	ρ_2	ρ_3	ρ_4	ρ_5	ρ_6
Panel B: Nominal ten-year spot rates, $y(10)$, and Real ten-year spot rates, $y^*(10)$										
$y(10)$	5.93	0.485	6.85	4.95	0.991	0.982	0.975	0.967	0.960	0.952
$y^*(10)$	3.87	0.310	4.42	3.28	0.996	0.991	0.987	0.983	0.979	0.974

Panel A provides sample statistics of the coupon STRIPS and TIPS datasets. The total number of coupon STRIPS security prices is 58,768, varying from 89 to 95 with an average of 93 securities per day. The total number of TIPS security prices is 4,390, varying from 6 to 8 with an average of 7 securities per day. Panel B reports sample statistics for estimated nominal and real ten-year spot rates. Mean, standard deviation (S.D.), maximum, minimum and first order till sixth order autocorrelations are presented, ten-year nominal spot rates, $y(10)$, are estimated using coupon STRIPS prices and the Bakshi, Madan and Zhang (2001) two-factor model. Ten-year real spot rates, $y^*(10)$, are adjusted for indexing lag effect based on estimates from TIPS prices using the single-factor CIR model (1985). In estimating ten-year nominal and real spot rates, we assume zero tax rates, zero transaction costs, and no arbitrage opportunities. Daily prices of TIPS and coupon STRIPS are collected from the Datastream database for the period from April 8, 1999, to September 7, 2001. In total, there are 632 observation days during the study period.

3. Data and estimation of nominal and real rates

The study uses daily U.S. Treasury STRIPS (stripped coupon only) and TIPS prices from the Datastream database. We also collect the reference CPI from the Bureau of Public Debt Online web site.[1] The study period covers April 8, 1999, through September 7, 2001. In total, there are 632 daily price observations. April 8, 1999, is the first date when there are at least six issues of TIPS outstanding. The study period ends on September 7,2001, to avoid the likely structure break in the estimated interest rate time series following the terrorist attack on September 11. The sample data include 58,768 STRIPS security prices and 4,390 TIPS security prices. Summary statistics for the datasets are reported in Table 1, Panel A.

We derive the nominal and TIPS term structures using daily trading prices of Treasury STRIPS and TIPS. The Cox, Ingersoll, and Ross (1985) (CIR) one-factor model is applied to estimate the term structure of interest rates for TIPS. Brown and Schaefer (1994) apply the CIR model to British government index-linked gilt prices. They find that the CIR model provides a flexible fit for the term structure of real interest rates. Once we obtain the TIPS term structure, the ten-year real spot rates are estimated by adjusting the lag effect of TIPS following Evans' (1998) method.[2]

The nominal term structure of interest rates is estimated using the two-factor framework derived by Bakshi, Madan and Zhang (2001). The two-factor model includes the

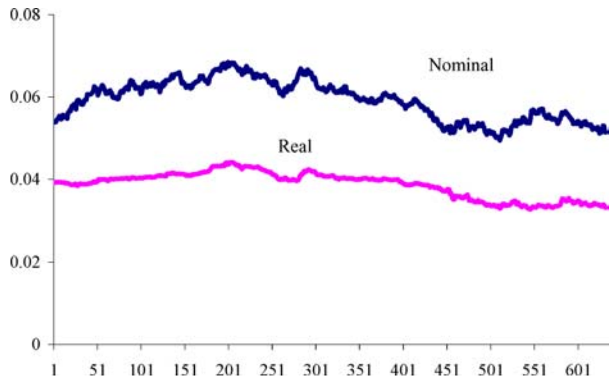

Figure 1. Real and nominal ten-year spot rates (Time period: April 8, 1999–September 7, 2001–Number of observations: 632) (*Notes*: Nominal rates are estimated from the two-factor model of interest rates. Real rates are first estimated from the one-factor CIR model and then adjusted for the indexing lag.)

instantaneous spot rate, and allows the long-run mean of the short-term rate to evolve stochastically. We prefer the two-factor model to the one-factor model for several reasons.

First, Bakshi, Madan and Zhang (2001) document that the application of a two-factor model significantly reduces both in-sample and out-of-sample mean-squared percentage pricing errors and absolute yield errors, compared to a one-factor model. Second, it has been suggested that returns on fixed-income securities can be explained by level, steepness, and curvature of the yield curve (Litterman and Scheinkman, 1991). Chen and Scott (1993) find that multi-factor models explain the changes in the steepness and the shape of the yield curve better than a one-factor model. Moreover, the two-factor model is preferable to the three-factor model because the third factor is not compatible with the real variability of bond prices (Chen and Scott, 1993).

Summary statistics for ten-year nominal, y(10), and real, $y^*(10)$, daily spot rates are presented in Table 1, Panel B.[3] The average ten-year nominal rate is 5.93% compared to a 3.87% real rate for the same time period. As expected, nominal rates also show a higher volatility than real rates. This higher volatility is further confirmed by examining the maximum and minimum values for these two spot rates series. High autocorrelations up to the sixth order are present for both ten-year nominal and real spot rates. The high autocorrelations indicate that both rates are nonstationary. Figure 1 shows the graphs of daily estimates of ten-year spot nominal and real rates between April 8, 1999, and September 7, 2001.

4. Methodology and results

4.1. Cointegration analysis

We estimate the integration order of the nominal and real rate time series using the augmented Dickey-Fuller (1979) (ADF) and Phillips-Perron (1988) (PP) unit root tests. The unit root test statistics are reported in Table 2 at 1 and 5% significance levels.

Table 2. Unit root test result for daily ten-year real and nominal spot rates (Time period: April 8, 1999-September 7, 2001—Number of observations: 632)

	Real rate			Nominal rate		
	Test Statistic	1%[a]	5%[a]	Test statistic	1%[a]	5%[a]
ADF[b]						
Level	0.15	−3.44	−2.87	−0.96	−3.44	−2.87
1st difference	−11.354[†]	−3.44	−2.87	−12.05[†]	−3.44	−2.87
PP[c]						
Level	0.15	−3.44	−2.87	−1.04	−3.44	−2.87
1st difference	−21.92[†]	−3.44	−2.87	−24.74[†]	−3.44	−2.87

[a]The critical value used here is available in MacKinnon (1991).
[b]The statistics are computed with one lag for the augmented Dickey-Fuller (ADF) test.
[c]The Phillips-Perron (PP) Test is performed with one non-zero autocovariance in the Newey-West (1987) correction for the heteroskedasticity and autocorrelation.
[†]The null hypothesis of a unit root is rejected at the 1% significance level.

Both the ADF and PP tests uniformly suggest that the ten-year real and nominal rate time series follow an $I(1)$ process. The hypothesis of the presence of a unit root cannot be rejected for the level series, while the presence of a unit root is rejected at the 1% significance level for the first-difference series.

The cointegration analysis tests for the presence of a long-term equilibrium relationship in two non-stationary interest rate series. The vector autoregression (VAR) model is characterized as follows:

$$\Delta Y_t = \mu + \Pi Y_{t-1} + \Gamma \Delta Y_{t-1} + \varepsilon_t \tag{1}$$

where ΔY_t, is a (2×1) first-difference time series vector; μ is a (2×1) constant vector; Π and Γ are (2×2) coefficient matrices; and ε_t is a (2×1) column vector of white Gaussian noise with mean zero and finite variance.[4]

The coefficient matrix Π incorporates information about the cointegration relationship among the variables in Y_t. Johansen (1988, 1991) demonstrates that the rank of the matrix Π reveals the number of cointegration relationships present among the variables in Y_t. Following the procedures developed by Johansen (1988), we can concentrate the model with respect to the matrix Π. The likelihood ratio test statistic for the hypothesis of at most r cointegration relationships and at least $m = 2 - r$ common trends is given by

$$\lambda_{\text{trace}} = -T \sum_{i=r+1}^{2} \ln(1 - \hat{\lambda}_i) \tag{2}$$

where T is the sample size actually used for estimation, and $\hat{\lambda}_1 > \hat{\lambda}_2$ are the eigenvalues of the squared canonical correlation between two residual vectors from level and first difference regressions. This is known as the Johansen trace test.

Table 3. Cointegration and Granger causality test results for daily real and nominal interest rates (Time period: April 8, 1999–September 7, 2001—Number of observations: 632)

Hypothesis	Eigenvalue	Likelihood ratio[a]	5 Percent critical value[b]	1 Percent critical value[b]
		Panel A: Johansen cointegration test		
$r = 0$	0.03365	21.95†	12.53	16.31
$r = 1$	0.00062	0.39	3.84	6.51

Null hypothesis	Real rate does not Granger cause nominal rate	Nominal rate does not Granger cause real rate
	Panel B: Granger causality test	
F statistic	0.28	4.86†
p-value	0.89	0.00

[a]The likelihood ratio test statistic is based upon the trace statistic: $\lambda_{\text{trace}} = -T\sum_{i=r+1}^{2}\ln(1-\hat{\lambda}_i)$, where $r = 0, 1$ and $\hat{\lambda}_i$ is the estimated i-th largest eigenvalue. λ_{trace} is the trace statistic for the hypothesis that at most there are r cointegrating relationships.
[b]The critical values for the trace statistic are available in Osterwald-Lenum (1992).
†Denotes rejection of the hypothesis at the 1% significance level.

The results of the Johansen cointegration tests are reported in Table 3 Panel A. The nominal and real rates are cointegrated with one cointegration relationship, indicating that there is one common stochastic trend. The trace test statistic of the null hypothesis that there is at most $r = 0$ cointegration vector is rejected at the 1% level, while the null of $r = 1$ cointegration vector is not rejected.

These results confirm that the two markets share one common long-run stochastic trend and one cointegration relationship. The time series of real and nominal rates do not drift too far apart over time. The result is consistent with our assumption at the outset that the TIPS market and the Treasury bond market form a cointegrated system.

4.2. Granger causality test

Johansen cointegration tests document that the ten-year real and nominal spot rates are cointegrated. While cointegration analysis is concerned with long-run equilibrium relationships, Granger causality tests are used to show short-term predictability of nominal and real spot rates. We apply Granger's (1969) approach to test whether the current real (nominal) rate can be explained by past values of the nominal (real) rate. A Granger causality relationship does not imply that one time series is the effect or the result of the other. Instead, it measures precedence and how information content flows between the two markets. The bivariate regression model is specified as follows:

$$y_t = \phi + \sum_{i=1}^{4}\varphi_i y_{t-i} + \sum_{i=1}^{4}\psi_i y^*_{t-i} + \varepsilon_t \tag{3}$$

$$y^*_t = \phi^* + \sum_{i=1}^{4}\varphi^*_i y^*_{t-i} + \sum_{i=1}^{4}\psi^*_i y_{t-i} + \varepsilon^*_t \tag{4}$$

The optimal length of lag is identified as four lags according to the Akaike information criterion. Wald statistics are calculated to test the joint hypothesis that $\psi_1 = \psi_2 = \psi_3 = \psi_4 = 0$ in Eq. (3) and $\psi_1^* = \psi_2^* = \psi_3^* = \psi_4^* = 0$ in Eq. (4). Test results are presented in Table 3 Panel B. We fail to reject the null hypothesis that the real rate does not Granger cause the nominal rate. However, the test statistic for the hypothesis that the nominal rate does not Granger cause the real rate is significant at 1% level. The results show that the real and nominal rates do not adjust simultaneously to new information; past nominal rates provide useful information to predict current real rates.

4.3. Common factor decomposition

Gonzalo and Granger's (1995) model is used to decompose the common factor in the cointegration system. This is done by decomposing an $I(1)$ series into permanent and transitory components, $I(1)$ and $I(0)$. Stock and Watson (1988) show that, for a cointegrated $I(1)$ series, there must exist a common factor representation in the form of:

$$Y_t = \theta f_t + \tilde{Y}_t \tag{5}$$

where θ is a loading matrix, f_t is a vector of an $I(1)$ common stochastic trend, and $\widetilde{Y}_t$, is a vector of $I(0)$ transitory component.

Following Johansen's framework, Gonzalo and Granger (1995) propose an estimation procedure of the common factor in a cointegrated system. The common factor is assumed to be driven by new information, and innovations to the common factor are permanent. The cointegration matrix Π in Eq. (1) has a reduced rank of $r < 2$, and can be decomposed as $\Pi = \alpha\beta'$, where α and β both are (2×1) matrices. The β matrix consists of the cointegration vectors, and α is the error correction (or equilibrium adjustment) matrix.

Gonzalo and Granger (1995) show that the common factor and the loading matrix are given by

$$f_t = \alpha_\perp' Y_t \tag{6}$$

$$\theta = \beta_\perp(\alpha_\perp'\beta_\perp)^{-1} \tag{7}$$

where $\alpha_\perp$ and $\beta_\perp$ are (2×1) matrices of full rank orthogonal to α and β, respectively. Under the hypothesis of cointegration, the maximum likelihood estimator of $\alpha_\perp$ can be found by solving the eigenvalues and eigenvectors of matrix $\hat{\Sigma}_{uu}^{-1}\hat{\Sigma}_{uv}\hat{\Sigma}_{vv}^{-1}\hat{\Sigma}_{vu}$. $\hat{\Sigma}_{\bullet\bullet}$ is the variance-covariance matrix for the residual vectors from the ordinary least squares regression. The subscripts u and v denote the level and first-difference regression, respectively. This will give the eigenvalues $\hat{\eta}_1 > \hat{\eta}_2$ and eigenvectors $\hat{M} = (\hat{m}_1, \hat{m}_2)$, normalized such that $\hat{M}'\hat{\Sigma}_{uu}\hat{M} = I$. The maximum likelihood estimator for the long-run common factor is then $\hat{\alpha}_\perp = (\hat{m}_{r+1}, \ldots \hat{m}_2)$, where r is the number of cointegration relationships in the system found using the Johansen approach. After $\alpha_\perp$ is estimated, the common factor f_t can be identified using Eq. (6). Specifically, the common permanent component is a linear combination of all variables in the cointegration system, where the weight for each variable is assigned by the matrix $\alpha_\perp$.

Table 4. Results of Gonzalo and Granger common factor model (Time period: April 8, 1999–September 7, 2001—Number of observations: 632)

Variable		Estimated coefficient for $\alpha_\perp$
Panel A: Maximum likelihood estimator of $\alpha_\perp$		
Nominal rate (y)		0.58
Real rate (y^*)		0.42
Null hypothesis	Real rate (y^*) not included in the common factor	Nominal rate (y) not included in the common factor
Panel B: Hypothesis tests		
Chi-squared statistic $\chi^2(1)$	1.80	11.39†
p-value	0.18	0.00

†Significant at the 1% level.

Gonzalo and Granger procedures are used to test whether any individual series alone or any combination of multiple series represents the common stochastic trend itself. First, according to the hypothesis to be tested, a (2×1) restriction matrix G is specified such that

$$H_0 : \alpha_\perp = G\theta \tag{8}$$

Second, the maximum likelihood ratio test of the hypothesis is obtained by first solving the eigenvalue $\hat{\Lambda}$ for $(G'\hat{\Sigma}_{uu}G)^{-1}G'\hat{\Sigma}_{uv}\hat{\Sigma}_{vv}^{-1}\hat{\Sigma}_{vu}G$, and then calculating the likelihood ratio as

$$-T \ln\left(\frac{1-\hat{\Lambda}}{1-\hat{\eta}_2}\right) \sim \mathcal{X}_1^2 \tag{9}$$

The estimated coefficient vector of the common factor, $\alpha_\perp$, is (0.58, 0.42)' as reported in Panel A of Table 4. The Chi-squared test for the TIPS market fails to reject the null hypothesis that the real rate series y_t^* is not included in the common factor. The Chi-squared statistic of 1.80 has one degree of freedom and a p-value of 0.18. The Chi-squared test does reject the null hypothesis that the nominal rate series y_t is not included in the common factor. It indicates that the Treasury bond market represents the dominant component of the common stochastic trend. The TIPS consistently defers to the Treasury bond market for its pricing.

4.4. *Vector error correction model*

We apply vector error correction models to further investigate the information flow between interest rate markets. If real rates respond to deviations from the long-run equilibrium relationship between real and nominal rates, but nominal rates do not respond to deviations between the two, we conclude that the information flows unilaterally from the Treasury bond market to the TIPS market.

The vector error correction model (VECM) is specified as:

$$\begin{aligned}\Delta y_t &= A\delta_{t-1} + \sum_{t=1}^{4} B_i \Delta y_{t-i} + \sum_{t=1}^{4} C_i \Delta y^*_{t-i} + \varepsilon_t \\ \Delta y^*_t &= a\delta_{t-1} + \sum_{t=1}^{4} b_i \Delta y_{t-i} + \sum_{t=1}^{4} c_i \Delta y^*_{t-i} + \omega_t\end{aligned} \tag{10}$$

where

$$\begin{aligned}\delta_t &= y_t - \kappa\, y^*_t \\ \Delta y_t &= y_t - y_{t-1} \\ \Delta y^*_t &= y^*_t - y^*_{t-1}\end{aligned}$$

δ_t represents the deviation of the real and nominal interest rates from their long-term equilibrium relationship specified by the cointegration coefficient κ.

The estimated cointegration coefficient κ is 1.54, as reported in Table 5. This long-term relationship reveals that for every 1% change in the expected real rate, there is a 1.54% change in the expected nominal rate. We would expect a positive relationship between the two variables because, according to Fisher (1930), the expected real rate is a component of every nominal rate. Fisher also theorizes that the expected inflation rate is a component of every nominal rate; however, we omit the expected inflation rate.

Evans (1998) documents a significant and positive coefficient between the expected real rate in Great Britain and the expected inflation rate as measured by Barclay's survey. The

Table 5. Results of vector error correction model (Time period: April 8, 1999–September 7, 2001—Number of observations: 632)

	$\Delta Y_t = A\delta_{t-1} + \sum_{i=1}^4 B_i \Delta y_{t-1} + \sum_{i=1}^4 C_i \Delta y^*_{t-1} + \varepsilon_t$ $\Delta y^*_t = a\delta_{t-1} + \sum_{i=1}^4 b_1 \Delta y_{t-i} + \sum_{i=1}^4 c_i \Delta y^*_{t-1} + \omega_t$ $\delta_t = y_t - 1.54y^*_t$, $\Delta y_t = y_t - y_{t-1}$, and $\Delta y^*_t = y^*_t - y^*_{t-1}$ Nominal rate Δ_{y_t}	Real rate $\Delta_{y^*_t}$
δ_{t-1}	−0.012(−1.07)	0.013(3.45)†
Δy_{t-1}	0.022(0.50)	0.036(2.28)†
Δy_{t-2}	−0.32(−0.72)	0.001(0.06)
Δy_{t-3}	−0.032(−0.71)	0.001(0.09)
Δy^*_{t-4}	0.010(0.22)	0.013(0.81)
Δy^*_{t-1}	−0.021(−0.17)	0.072(1.63)
Δy^*_{t-2}	0.016(0.13)	−0.011(−0.24)
Δy^*_{t-3}	−0.015(−0.12)	−0.075(−1.70)
Δy^*_{t-4}	−0.018(−0.14)	−0.011(−0.24)

t-statistics in parentheses.
†indicates the test statistics are significant at 1% level.

regression coefficient for the expected real rate ranges from 0.61 to 0.82. Evans' finding is consistent with the coefficient of 1.54 that we find with an omitted expected inflation variable. The coefficient of 1.54 for the real rate suggests a combination of a one-to-one relationship with the expected nominal rate and a 0.54 indirect relationship with expected inflation.

We also test the statistical significance of the coefficient terms, A *and* a for δ_{t-1}. A significant coefficient indicates that the adjustment process reflects the divergence from the long-term equilibrium relationship between real and nominal interest rates. A significant and positive estimated coefficient $\hat{a}$ suggests that the real interest rate adjusts itself according to the path of the nominal rate. If the previous nominal rate is high compared to the real rate, the real rate tends to adjust upward in the next period to reduce the deviation. Meanwhile, if the estimated coefficient $\hat{A}$ is not significantly different from zero, we conclude that the nominal rate follows its own path and does not defer to the TIPS market for its pricing.

Table 5 shows that the error correction term in the nominal rate equation is insignificant, but it is significant in the real rate equation. The t-statistic of the error correction term in the real rate equation is 3.45. This suggests that the real rate series adjusts to the error correction terms in the system, while the nominal series does not. Through an error correction mechanism, the real rate keeps itself from moving too far away from the long-run equilibrium relationship. A tendency to align with the equilibrium relationship does not appear in the nominal rate series. This evidence implies that the nominal Treasury bond market appropriately impounds the common stochastic factor. Table 5 also reports the regression coefficients for lagged variables. Except for the coefficient related to the first lagged change of the nominal rate in the real rate equation, none of them is significant at the 5% level. The coefficients indicate that the error correction process takes no more than one additional day to accomplish. The results highlight the degree of information risk in the TIPS market.

5. Summary and conclusions

This study investigates the presence of information risk in TIPS market. The information flows between the nominal Treasury bond market and the TIPS market determine nominal and real interest rates in the U.S. Our motivation is to document whether information risk as described by O'Hara (2003) is a plausible explanation of the surprisingly high real yield of the TIPS market during its trading history. Information risk exists when the flow of new information relevant to two markets occurs more quickly in one market than in another market, leaving the second market disadvantaged. While traders in the TIPS market are attracted to a security offering inflation protection, they are disadvantaged in information risk and will demand compensation for incurring the risk.

The linkage between the two markets is explained by Fisher's theory of interest rate determination. The equilibrium spread between real and nominal rates should not differ from the sum of expected inflation and an inflation risk premium. Applying unit root tests to ex ante nominal and ex ante real ten-year spot rates provides evidence that both time series follow an $I(1)$ process. There is only one cointegrating relationship between the nominal and real interest rates, according to Johansen's cointegration analysis.

The Granger causality test indicates that the nominal and real interest rates do not respond simultaneously to new information and that the nominal rate Granger causes the real rate. The Gonzalo and Granger common-factor model (1995) indicates that the nominal bond market is the dominant market for price innovation in these two markets. Chi-squared statistics fail to reject the null hypothesis that the real rate is not included in the common stochastic trend. The VECM analysis indicates that information flows unilaterally from the nominal bond market to the TIPS market. The unilateral information flow between less transparent bond markets is distinctly different from previous empirical findings of bilateral or multilateral information flows among more transparent equity and futures markets.

The two interest rate markets respond differently to deviations from their equilibrium relationship. Price innovation from a random shock occurs primarily in the nominal rate market. The nominal rate follows its own path and does not defer to the TIPS market for its pricing. The TIPS market, as a follower, takes one additional day to adjust its equilibrium relationship with nominal rates. The TIPS market bears the responsibility for convergence to the equilibrium spread, after the new information is reflected initially in the nominal Treasury market.

The findings are consistent with a nominal bond market that reveals new information about U.S. nominal and real interest rates more quickly than the TIPS market, and are consistent with our expectation that informed traders concentrate their trades in the established nominal Treasury markets. If TIPS investors suspect that they are subject to information risk, then they would be expected to demand higher compensation for bearing this risk. Our empirical finding of a one-day lag quantifies the degree of information risk for TIPS investors, who may now adjust their current perceptions of the extent of their informational disadvantage. This knowledge may prompt additional studies of ways to further improve the flow of information to the TIPS market such as through the trading of TIPS derivatives. Finally, a future extension of the paper is to track the information risk in the TIPS market beyond the nascent trading period covered by this study.

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Notes

1. The address for the website is: http://www.publicdebt.treas.gov.
2. For detailed discussion of estimation processes of the nominal and real interest rate term structures, as well as CPI indexing procedure for TIPS, please see Chu, Pittman and Yu (2003).
3. Parameters estimates for the daily nominal and TIPS term structures are available upon request. Chu, Pittaman and Yu (2003) reports parameters estimates for weekly nominal and TIPS term structures over the time period from April 14, 1999 through April 4, 2001.
4. In our study, the Y_t vector is the transpose of $[y_t(10), y_t^*(10)]$.

References

Bakshi, G., D. Madan, and F. Zhang, "Investigating the Sources of Default Risk: Lessons from Empirically Evaluating Credit Risk Models." Working Paper, The Federal Reserve Board, Finance and Economics Discussion Series, 2001.

Benninga, S. and A. Protopapadakis, "Real and Nominal Interest Rates under Uncertainty: The Fisher Theorem and the Term Structure." *Journal of Political Economy* 91, 856–867 (1987).

Brown, R. H. and S. M. Schaefer, "The Term Structure of Real Interest Rates and the Cox, Ingersoll, and Ross Model." *Journal of Financial Economics* 35, 3–42 (1994).

Chatrath, A., M. Chaudhry, and R. Christie-David, "Price Discovery in Strategically Linked Markets: The TED Spread and Its Constituents." *Journal of Derivatives* 9, 77–87 (1999).

Chen, R. and L. Scott, "Maximum Likelihood Estimation for a Multifactor Equilibrium Model of the Term Structure of Interest Rates." *Journal of Fixed Income* 3, 14–31 (1993).

Chu, Q. C., D. N. Pittman, and L. Q. Yu, "Real Rates, Nominal Rates, and the Fisherian Link." *International Review of Financial Analysis* 12, 189–205 (2003).

Chu, Q. C., D. N. Pittman, and L. Q. Yu, "Information Content of Maturing TIIS." *Journal of Fixed Income* 13, 90–99 (2004).

Cox, J. C., J. E. Ingersoll, and S. A. Ross, "A Theory of the Term Structure of Interest Rates." *Econometrica* 53, 385–408 (1985).

Dickey, D. A. and W. A. Fuller, "Distribution of the Estimators for Autoregressive Time Series with a Unit Root." *Journal of the American Statistical Association* 74, 427–431 (1979).

Evans, M. D. D., "Real Rates, Expected Inflation, and Inflation Risk Premia." *Journal of Finance* 53, 187–218 (1998).

Easley, D. and M. O'Hara, "Information and the Cost of Capital." *Journal of Finance* 59, 1553–1583 (2004).

Fisher, I., *The Theory of Interest*. New York: Macmillan, 1930.

Gonzalo, J. and C. W. J. Granger, "Estimation of Common Long-Memory Components in Cointegrated Systems." *Journal of Business & Economic Statistics* 13, 27–35 (1995).

Granger, C. W. J., "Investigating Causal Relations by Econometric Models and Cross Spectral Methods." *Econometrica* 37, 424–438 (1969).

Grossman, S. J. and J. Stiglitz, "On the Impossibility of Informationally Efficient Markets." *American Economic Review* 70, 393–408 (1980).

Harris, F. H., T. H. McInish, and R. A. Wood, "Security Price Adjustment Across Exchanges: An Investigation of Common Factor Components for Dow Stocks." *Journal of Financial Markets* 5, 277–308 (2002).

Johansen, S., "Statistical Analysis of Cointegration Vectors." *Journal of Economic Dynamics and Control* 12, 231–254 (1988).

Johansen, S., "Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models." *Econometrica* 59, 1151–1580 (1991).

Laatch, F. E. and D. P. Klein, "Nominal Rates, Real Rates and Expected Inflation: Results from a Study of U.S. Treasury Inflation-Protected Securities." *Quarterly Review of Economics and Finance* 43, 405–417 (2003).

Litterman, R. and J. Scheinkman, "Common Factors Affecting Bond Returns." *Journal of Fixed Income* 1, 54–61 (1991).

MacKinnon, J. G., "Critical Values for Cointegration Tests." In *Long-Run Economic Relationships*, R. F. Engle and C. W. J. Granger (Eds.), Oxford University Press, 1991, pp. 267–276.

Newey, W. K. and K. D. West, "A Simple Positive Semi-Definite, Heteroskedasticity and Autocorrelation Consistent Covariance Matrix." *Econometrica* 55, 703–708 (1987).

O'Hara, M., "Presidential Address: Liquidity and Price Discovery." *Journal of Finance* 58, 1335–1354 (2003).

Osterwald-Lenum, M., "A Note with Quantiles of the Asymptotic Distribution of the Maximum Likelihood Cointegration Rank Test Statistics." *Oxford Bulletin of Economics and Statistics* 54, 461–472 (1992).

Phillips, P. C. B. and P. Perron, "Testing for a Unit Root in Time Series Regression." *Biometrika* 75, 335–346 (1988).

Roll, R., “Empirical TIPS.” *Financial Analysts Journal* 60, 31–53 (2004).

Sack, B., “Deriving Inflation Expectations from Nominal Inflation-Indexed Treasury Yields.” *Journal of Fixed Income* 10, 6–17 (2000).

Sack, B. and R. Elsasser, “Treasury Inflation-Indexed Debt: A Review of the U.S. Experience.” *Economic Policy Review* 10, 47–63 (2004).

Shen, P. and J. Coming, “Can TIPS Help Identify Long-Term Inflation Expectations?” *Economic Review*, Federal Reserve Bank of Kansas City 86, 61–87 (2001).

Stock, J. H. and M. W. Watson, “Testing for Common Trends.” *Journal of the American Statistical Association* 83, 1097–1107 (1988).

Umlauf, S. R., “Information Asymmetries and Security Market Design: An Empirical Study of the Secondary Market for U.S. Government Securities.” *Journal of Finance* 46, 929–953 (1991).