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Arbitrage in International Sovereign Debt Markets? Evidence from the Inflation-Protected Securities of Six Countries

We consider an arbitrage strategy that exactly replicates the cash flow of a sovereign nominal bond using inflation swaps and inflation-linked bonds. The strategy reveals a violation of the law of one price in the G7 countries, which is largest for the eurozone. Testing the strategy's exposure to deflation, volatility, liquidity, and macro-economic risks shows the observed mispricing is a risk premium, which is more pronounced in the eurozone. We find less support that financial limits to arbitrage explain the mispricing. We conclude that pure long-run arbitrage opportunities persist when these strategies are exposed to intermediate financial risks.

JEL codes: G12, G15, G18, H63
Keywords: inflation-indexed bonds, law of one price, limits to arbitrage, mispricing, nominal bonds

As noted by Fleckenstein, Longstaff, and Lusting (2014), there is a substantial violation of the law of one price across the treasury nominal and inflation-protected security markets. This violation occurs because there is a difference in price (which, consistent with the finance literature, we call the mispricing) between a sovereign nominal bond and a synthetic bond, which replicates the nominal bond's cash flows using inflation-protected securities and inflation swaps. As

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an apparently substantial arbitrage opportunity in some of the most prominent and liquid financial markets in the world, this mispricing provides a direct challenge to no-arbitrage conditions common in financial market theory.

In this paper, we provide new evidence on this violation of the law of one price for a large number of nominal and inflation-protected securities in the United States, Europe, and Japan. We then examine the drivers of mispricing and test for the limits of arbitrage as a potential explanation for the mispricing. We examine the arbitrage strategy's exposure to deflation risk, volatility risk, liquidity risk, and macroeconomic risk. We also examine the sensitivity of the mispricing to the funding costs of arbitrageurs. We look at this large array of factors because one explanation for this mispricing would be that inflation-protected securities pay higher yields because of an increased exposure to a risk factor like liquidity risk. However, the divergence also leads to a substantial arbitrage opportunity possibly due to market segmentation and the limits of arbitrage to reduce this mispricing.

First, we apply a replication strategy where we use the market prices of inflation swaps and sovereign inflation-indexed bonds to derive synthetic nominal bonds that replicate exactly the cash flow of sovereign nominal bonds in six of the seven G7 countries. The first part of our analysis closely follows Fleckenstein, Longstaff, and Lusting (2014), who apply a similar arbitrage strategy for the U.S. linkers. Our analysis spans the period February 02, 2007, to November 30, 2012. Our analysis includes 25 matches for the United States, 5 matches for the United Kingdom, 3 matches for Japan, 4 matches for Germany, 5 matches for France, and 5 matches for Italy, yielding a total of 47 bond pairs. We obtain our data from the Bloomberg system.

We find evidence of a pricing anomaly that is substantial for most securities in all the countries. On average the synthetic bond, which perfectly replicates the cash flow of the nominal bond, is cheaper than the nominal bond itself. The average pricing anomaly in the sample of U.S. nominal bonds is \$1.67, less than the figure of \$3.13 reported by Fleckenstein, Longstaff, and Lusting (2014) using data for an earlier period.[1] The reduction in the magnitude of the average mispricing might imply that the pricing anomaly has diminished with time, as the amount of capital available to arbitrageurs increases. We examine this conjecture later in this paper. An alternative notion that we investigate is that the risk factors to which the mispricing is exposed, for example, the possibility of an extended period of low economic growth and deflation, have subsided with the settling of the financial crisis and the now more normal functioning of financial markets. The lowest average pricing anomaly in our sample is \$1 for France followed by Japan with mispricing of \$1.74 and the United Kingdom \$1.93. Italy displays the largest pricing anomaly of \$8.71 followed by Germany with \$3.12.

We then examine the factors, which correlate with this mispricing. We find that the mispricing is well modeled as a compensation for risk. Specifically, the arbitrage strategy appears to be exposed to volatility risk (as measured by the volatility index [*VIX*]) and deflation risk (as measured by inflation risk premia [*IRPs*]).

1. We follow the literature and express the mispricing in U.S. dollars per 100 notional.

This result is due to the fact that the less liquid treasury inflation–protected security (TIPS) require compensation to be held in these states. On the other hand, we find little evidence that when arbitrageurs have more capital that the measured mispricing narrows. This result suggests that limits to arbitrage does not solely explain the mispricing. Next, using a structural vector autoregression (VAR), we consider the reaction of the mispricing to an unexpected change in the short-term interest rate. We find that this change reduces the mispricing in the short run but increases it in the long run. This result gives more credence to the risk explanation of the mispricing because one would think a pure arbitrage opportunity would widen as the cost of funding to arbitrageurs increases.

Once we establish the correlation of the mispricing with risk factors, we then treat the eurozone crisis as an ideal environment to investigate these risk factors in more detail. During the sample period, relative to the United States, United Kingdom, and Japan, the eurozone was exposed to more economic risk: for example, default, deflation, and downside economic risk. Take, for example, Italy—with the largest mispricing of \$8.71—a eurozone country whose credit rating was downgraded by Moody' s on October 4, 2011, from Aa2 to A2, and by the end of the sample period on July 13, 2012, had a further downgrade to Baa2 owing to the size of its public debt. On average, we find the eurozone countries to have over two times larger mispricing than the non-euro countries. The average mispricing for the eurozone countries is \$3.95 while the non-euro countries display a mispricing of \$1.67. Using a difference-in-difference regression we find that the mispricing increased by \$3.40 relative to the non-euro countries during the eurozone crisis. Additionally, the mispricing is more highly correlated with the risk factors we isolate and the magnitude of the coefficients are larger, allowing us to conjecture that the pricing anomaly reported in this paper is an economic tail risk premium rather than an easily exploitable arbitrage opportunity.

While the mispricing variable we calculate is not a measure of *inflation expectations*, our paper relates to a large literature in macroeconomics, which pursues the measurement of *inflation expectations*. This literature is extremely valuable as accurate and timely measures of *inflation expectations* are of great value to monetary policymakers. One common measure of *inflation expectations* is the break-even inflation rate, the difference between the yield on a treasury bond and the yield on a TIPS of the same maturity. However, it is well known that as a pure measure of *inflation expectations* this break-even inflation rate suffers from many problems. For example, the different risk exposures on nominal and inflation-protected debt, and the time varying impact of these risk factors, prevents the break-even inflation rate from being a pure measure of *inflation expectations*. To address these problems, there is a substantial literature (e.g., Christensen et al. 2010), which backs out break-even inflation rates from Treasury and TIPS data by modeling the mispricing as a compensation for risk. In this paper we find that the mispricing is, at least in part, a compensation for bearing risk. As such, it lends credence to these structural models, which measure inflation expectation from structural risk models of the term structure.

The remainder of this paper is organized as follows. Section 1 reviews the relevant literature. Section 2 describes the replicating strategy. Section 3 describes the data. Section 4 discusses the econometric strategy and results. Section 5 examines the results for the eurozone and Section 6 concludes.

1. LITERATURE REVIEW

The zero-arbitrage relationship between the U.S. Treasury inflation–indexed bonds (TIPS) and nominal treasury bonds was originally analyzed by Fleckenstein, Longstaff, and Lusting (2014). Later studies by Haubrich, Pennacchi, and Ritchken (2012) and Fleckenstein (2013) confirm their key findings. In this literature, mispricing is attributed primarily to investors' preferences for the safety and liquidity of nominal treasury bonds (Longstaff 2004, Bansal, Coleman, and Lundblad 2010). Our results corroborate the findings of Fleckenstein, Longstaff, and Lusting and Fleckenstein that the convenience yield attributed to nominal treasury bonds does not extend to inflation-indexed bonds and therefore there is a substantial violation of the law of one price across the nominal and inflation-linked bond markets.

However, our work differs from Fleckenstein, Longstaff, and Lusting (2014) in several respects. First, our analysis extends to international markets by including six of the G7 countries and extends the sample period through 2012 to include the eurozone crisis period.[2] We also consider a relatively large sample of 47 pairs of bonds. Further, our analysis is at the individual security level rather than in aggregate to avoid any possible systematic patterns that can influence the pricing anomaly if analyzed in aggregate. Additionally, we focus on the arbitrage strategy's exposure to deflation, volatility, liquidity, and macro-economic risks. We also complement these studies by including the *IRP* as an additional important risk factor. Further, we explicitly distinguish between the postfinancial and euro-crisis period as an ideal environment to study the deflationary pressures in the eurozone with respect to the rest of the market.

This paper also contributes to the literature on the persistence of mispricing and arbitrage opportunities. Gromb and Vayanos (2002) and Ashcraft, Gârleanu, and Pedersen (2011) show that margins, haircuts, and other frictions may induce deviations from the law of one price. Brunnermeier and Pedersen (2009) examine the effect of liquidity on security prices. Duffie (2010) examines the relationship between slow-moving capital and mispricing in financial markets. Deviations from the law of one-price have been rationalized in the literature in several ways, including liquidity effects, liquidity risk premia, and arbitrage risk premia. Haubrich, Pennacchi, and Ritchken (2012) and Christensen and Gillan (2011) characterize the component of the inflation-indexed bond price that cannot be explained using a formal asset pricing

2. D'Amico et al. (2016) and Grishchenko and Huang (2013) on the other hand do not include data beyond March 2007, similarly Fleckenstein et al. (2014) spans through November 2009. Gürkaynak and Wright (2012) document significant pricing discrepancies with comparable maturity bonds trading at quite different prices in November and December of 2008.

model as a liquidity risk premium. We test the predictions of the slow-moving capital theory by examining the relationship between the change in the capital available to arbitrageurs and the levels and differences of a mispricing measure as well as the exposure of the arbitrage strategy to various risk factors.

While the *IRP* is not a central focus of our paper, we explore inflation risk as a potential driver of the mispricing. There is a large literature attempting to measure the interest rate premium for bearing inflation risk. See, for example, Campbell and Shiller (1996), Campbell and Viceira (2001), Buraschi and Jiltsov (2006), and Ang, Bekaert, and Wei (2008). In our paper, we follow the approach of Haubrich, Pennacchi, and Ritchken (2012) and use data on the inflation swap market. Specifically, we measure an *IRP* as the difference between inflation swap rates and survey measures of expected inflation. We then see if the mispricing is correlated with this inflation risk.

Our paper also relates to the extensive literature, which extracts *inflation expectations* from inflation linked debt, for example, Christensen, Lopez, and Rudebusch (2010), D'Amico, Kim, and Wei (2016), Hördahl and Tristani (2012), Adrian and Wu (2010). These papers argue that the difference in yields on nominal and inflation-protected debt are influenced both by expected inflation and risk factors, for example, liquidity risk. Our paper is generally consistent with this view, as we find that the mispricing in the nominal and inflation linked debt markets is correlated with our proxies for risk factors.

2. ARBITRAGE STRATEGY

The arbitrage strategy that we follow has been long recognized and applied by practitioners.[3] The essence of this strategy is to replicate the cash flow from a nominal sovereign bond by using a synthetic bond formed from a combination of inflation-protected bonds, inflation swaps, and zero coupon bonds (strips). An arbitrageur then purchases the cheaper bond and sells the more expensive bond. She locks in a positive profit today with zero net cost in the future.

To explain this strategy in more detail, Figure A1 in the Appendix, represents the approach graphically. A nominal bond entitles the bearer to receive a cash flow $CF_t^{Nominal}$ at time t. This cash flow can be replicated by first purchasing an inflation linked bond, which promises, at time t, a cash flow of $CF_t^{Tips} I_t$ where I_t is the gross inflation rate $\frac{P_t}{P_0}$ between now and time t. The arbitrageur then enters into an inflation swap, which allows her to exchange that future cash flow $CF_t^{Tips} I_t$ for a nominal cash flow, CF_t^{Swap}, which is known today. Finally the arbitrageur purchases zero-coupon bonds today to guarantee a known cash flow CF_t^{Strips} at time t so that $CF_t^{Strips} + CF_t^{Swap} = CF_t^{Nominal}$. Therefore, the arbitrageur has created a synthetic bond (made from inflation-protected bonds, swaps and zero coupon bonds) that

3. See *Financial Times* blog of April 4, 2012, *Wall Street Journal* April 27, 2010, among others, that discuss this strategy.

replicates the payoffs of the nominal bond. The final step is to buy the cheaper bond today, sell the more expensive one, and lock in a positive profit with zero net cost in the future. Based on this strategy then, we can define our main variable of interest, mispricing, as the price of the nominal bond $P_{nominal}$ minus the price of the synthetic bond P_{syn}.

The arbitrage strategy is executed in the same way for all six countries included in the study. The number of days between the maturity of each inflation-indexed bond and the nominal bond with the nearest maturity is defined as maturity mismatch. To adjust for maturity mismatch, the yield to maturity on the synthetic bond is applied to obtain the price of a hypothetical synthetic bond that would match precisely the maturity of the nominal Treasury bond in the pair. For any maturity mismatch, the cash flows of the synthetic bond always match those of the underlying nominal bond precisely, by construction. The mispricing is analyzed for each security individually to avoid any possible systematic patterns that can influence the mispricing if analyzed in aggregate.

3. DATA

The data comprise daily closing prices for sovereign government nominal bonds, government inflation–indexed bonds, strips and inflation swaps for six countries: the United States, United Kingdom, Japan, Germany, France, and Italy. The observation period is February 2, 2007, to November 30, 2012, for the majority of the securities analyzed.[4] We obtain the data from Bloomberg. The inflation-indexed bonds and nominal bonds have various maturities from 2008 to 2032. The nominal and inflation-indexed bond daily prices are adjusted for accrued interest, following the standard conventions.

Inflation swaps are quoted in terms of a constant rate on the contract's fixed leg. The traded maturities are 1, 2, 3, 4, 5, 6, 7, 8, 9, 10, 12, 15, 20, 25, and 30 years for the United States, United Kingdom, and Japan. For the eurozone countries, the longest maturity for an inflation swap is 25 years. We interpolate for intermediate swap maturities.

We match the inflation-indexed bonds and nominal bonds as closely as possible, based on their respective maturities. The maturity mismatch is defined as the number of days between the maturity of the inflation-indexed bond, and the maturity of the nominal bond with the closest maturity. We examine all sovereign inflation-indexed nominal bond issues for six of G7 countries available on Bloomberg system for the time period analyzed yielding 25 pairs of bonds for the United States, 5 pairs for the

4. For some countries we have a slightly shorter time period. Japan begins in February 2008. Germany begins in May 2008. In addition for some bond pairs we have shorter time periods, the shortest being March 21, 2012, to December 6, 2012, for a pair from Germany or July 24, 2011, to November 30, 2012, for a pair from the United States.

United Kingdom, 3 pairs for Japan, 4 pairs for Germany, 5 pairs for France, and 5 pairs for Italy.

In addition to the bond market data used to calculate the mispricing, we use several variables to examine whether the observed pricing anomaly correlates with financial or macro-economic variables. This analysis is important because although the arbitrage strategy is profitable if held to maturity, a widening of the mispricing may require an arbitrageur to liquidate the strategy prematurely, incurring significant losses (see Shleifer and Vishny 1997). For example, if inflation, particularly anticipated inflation, induces a rapid reduction in the value of the underlying assets, this effect would reduce the arbitrageurs' engagement in this trading strategy. Further, the relation of financial and macrovariables with the observed pricing anomaly would also reveal important information on the market' s assessment of deflation risk and other relevant economic tail risks.

The following variables are all obtained from the Bloomberg system for each country. The first variable we use is the 10-*year swap spread*, as a principal proxy for the credit risk of the banking system. Next, we obtain sovereign Credit Default Swap (CDS) spreads for each country in our analysis. CDS spreads should capture all relevant information concerning the altered risk of default for each country. Since CDS insures holders against any financial losses resulting from a credit event, it provides a quantitative measure of the risk associated with sovereign debt. Additionally, we use CDS prices to examine the extent to which sovereign default risk contributes to the mispricing. These portfolios will provide information on the extent to which default risk contributes to the mispricing. Finally we collected the *VIX* (an option-implied volatility index) for the stock market of each country. *VIX* is widely considered as the "fear index" since it reflects market's assessment of the risk of a large downward movement in the stock market, an interpretation we will use in our subsequent results.

Next, we collect data on the *inflation risk premium* (*IRP*)s. We use market participant's conventional definition of *IRP*, the difference between the inflation swaps and expected inflation rates. Higher inflation swap rate than the expected inflation rate implies positive *IRP* and vice versa. Since there is no theoretical reason for *IRP* not to be negative, the occurrence of this scenario can be therefore viewed as a deflation risk premium.[5]

We also examine the impact of several macro-economic variables on the mispricing variable. We are particularly interested in the ability of macro-economic variables to explain time series variation in the mispricing and to capture realized macroeconomic risk over time. These variables are *oil prices*, *overnight bank lending rates*, *industrial production*, *government deficits*, and *inflation expectations*. *Oil prices* are

5. To measure *inflation expectations* we take data from the University of Michigan survey data for the United States; Bank of England Survey of External forecasters; Bank of Japan Inflation Outlook of Enterprises (Tankan) for the Japan; and European Central Bank (ECB) inflation forecasts for the eurozone countries. University of Michigan data can be accessed from the Federal Reserve Bank of Philadelphia website https://www.philadelphiafed.org/research-and-data/real-time-center/survey-of-professional-forecasters/; Bank of England website http://www.bankofengland.co.uk/research/Pages/onebank/datasets.aspx; Tankan is available at https://boj.or.jp/en/statistics/tk/index.htm; ECB data are available at https://www.ecb.europa.eu/stats/prices/indic/forecast/html/table_hist_hicp.en.html.

crude oil (West Texas Intermediate) spot prices. We use *oil prices* to capture the state of the global economy. Given that during the time period of this study *oil prices* tended to rise on good economic news, higher *oil prices* should be associated with improved expectations of economic conditions. *Overnight bank lending rates* are the Fed Funds rate from Federal Reserve Economic Database (FRED) for the United States, for Japan, it is the basic loan rate, for the United Kingdom and the euro area countries, we use Libor and Euribor. We use the overnight bank lending rate as a potential measure of the cost of funding for banks and other financial institutions investing in the bond markets. *Industrial production* is used because it is available monthly and gives an indication as to the state of the economy. *Inflation expectations* are used here because increased *inflation expectations* could increase demand for inflation-protected securities and also be consistent with an improved outlook on the economy. Finally, we examine the role of *government deficits* as they can be associated with larger default risk and potentially affect bond prices.[6]

Lastly, we are also interested in the role arbitrageurs play in reducing the mispricing. To that end we collect data from Bloomberg system on the *HFRX hedge fund index returns*. As subcategories we examine the HFRX macrostrategy index, relative value hedge fund index, the all fixed-income convertible arbitrage index, the fixed-income sovereign index, and the global index returns. We choose these hedge fund categories because they are the hedge funds most likely to engage in the type of arbitrage strategy that would reduce the mispricing. We have also explored the role that supply of bonds—defined as new issuance of nominal debt and inflation-linked debt relative to total government debt—as an additional institutional factor.[7] However, we have not found it to be significant in the regressions so have omitted the results. For clarity, we have gathered these variables in Table A1 in the Appendix along with the relevant Bloomberg tickers.

3.1 The Inflation-Indexed Bond and the Inflation Swap Markets

Inflation-indexed bonds are direct obligations of sovereign states. The key difference between a nominal sovereign bond and an inflation-indexed bond is that the principal of the latter is adjusted over time to reflect changes in the Consumer Price Index (CPI). The fixed coupon rate for the inflation-indexed bond is applied to the principal, so the (semi)annual coupon payment varies in line with the adjustment to the principal for realized inflation or deflation. The repayment of principal at

6. Data on *oil prices*, the Fed Funds rate, and *industrial production* come from the FRED. Overnight bank rates for Japan, United Kingdom, and the euro countries come from Bloomberg. Data on *inflation expectations* come from the Survey of Professional Forecasters administered by the Federal Reserve Bank of Philadelphia for the United States and the ECB for the euro area countries. Data are available at: https://www.philadelphiafed.org/research-and-data/real-time-center/survey-of-professional-forecasters/data-files and http://www.ecb.europa.eu/stats/prices/indic/forecast/html/index.en.html, respectively. Finally, data on *government deficits* come from the Organization for Economic Co-operation and Development (OECD), https://data.oecd.org/gga/general-government-deficit.htm.

7. Following Robin and Dimitri (2014) and Pflueger and Viceira (2011), supply is defined as Supply $= D^{TIPS}/D_t$, where D^{TIPS} is the face value of the outstanding inflation-indexed bonds and D_t is the total government debt. Change in supply is defined as $\triangle Supply_t = (D_t^{TIPS} - D_{t-1}^{TIPS})/D_{t-1}^{TIPS} - (D_t - D_{t-1})/D_{t-1}$.

maturity is the higher of the original principal or the inflation-adjusted principal. While the coupon payments can be reduced in the event of deflation, the same does not apply to the repayment of principal, this is known as redemption floor or deflation protection.

While most issuers guarantee the principal floor of 100, not all issuers have this deflation protection option. In our sample, only the UK and Japanese linkers do not have an embedded deflation protection option. The deflation floor is effectively a put option whose value is influenced by the movements of the relevant price index from the time of issuance. A spike in the inflation or CPI moves the option value out of the money rendering the newly issued bonds of the same issuer more attractive than the seasoned ones, even when all things remain equal since the newly released bonds are in the money. The presence of a deflation floor tends to make the mispricing less correlated with deflation risk. This is because in general the nominal bonds are more expensive than the synthetic bonds. As the economy moves closer to a deflation scenario, the value of the deflation option rises, and the mispricing will narrow, creating a negative correlation between deflation risk and the mispricing. Therefore, we do not believe that the deflation option drives the correlations we find with financial and macro-economic risks.

Another important cross-country variation in the security design is the indexation lag and the coupon frequency. Inflation-linked bonds are usually referenced to a relevant domestic cost-of-living index. For example, U.S. TIPS are indexed to the nonseasonally adjusted CPI for All Urban Consumers (the CPI-U). In the United Kingdom for older issues this was the Retail Price Index (RPI) and for more recent United Kingdom issues the CPI. In the euro area, this is the weighted average of the individual euro area countries' harmonized price indices, the Harmonised Index of Consumer Prices excluding tobacco (HICP ex-tobacco). While the majority of the linkers studied in this paper have a 3-month lag indexation, the UK inflation-linked (IL) gilts for issues prior to 2005 had an indexation lag of 8 months. After 2005 United Kingdom' s (IL) gilts have a 3-month indexation lag, as do all other linkers analyzed in this paper. A final distinction in the security design among heterogeneous sovereign indexed bonds issuers is the frequency at which coupons are paid. Except for Germany and France who pay coupons annually, the United States, the United Kingdom, Italy, and Japan pay coupons semiannually. The indexation lag and the coupon payment frequency affect how well the linkers compensate for contemporaneous inflation. For robustness, we have rerun our main results (Table 3), exclude the UK gilts due to the higher indexation lag and the lack of deflation protection floor and Japan's bonds as well. We then also excluded the German and French pairs due to their annual, as opposed to the semiannual, coupon issuance. These changes did not affect our main results.

3.2 Size of the Swap Market and Taxation of Bond Income

While there are not many reports on the size of the inflation swap markets, Fleming and Sporn (2013), using novel transaction data, estimate the bid–ask spreads of the U.S. inflation swap market to be in the region of two-to-three bps and an

TABLE 1

Summary Statistics of Mispricing

	Mean	Median	SD	Min	Max	*N*
United States	1.668	1.314	5.979	–20.95	35.11	911
United Kingdom	1.929	0.515	6.701	–14.49	25.26	229
Japan	1.738	1.567	2.343	–5.842	9.621	93
Germany	3.121	2.729	2.013	–0.344	8.393	106
France	1.008	1.051	2.775	–6.432	10.50	226
Italy	8.712	4.740	9.484	–1.132	31.92	255
Europe	4.736	2.557	7.442	–6.432	31.92	587
Non-Europe	1.722	1.286	5.927	–20.95	35.11	1233

Note: This table reports the summary statistics for the nominal bond minus synthetic nominal bond mispricing for the 47 pairs of six of G7 countries. The mispricing is measured in dollars per $100 notional. Mean, Median, Std Dev, Min, Max, and *N* report the average and median dollar mispricing for each pair we analyze, its standard deviation, the highest and lowest mispricing values, and the number of monthly estimations for each security for each country, respectively. The sample period spans from February 2, 2007, to November 30, 2012.

average transaction size of $65 million per day. Furthermore, the authors show that the average difference between the Bloomberg quoted inflation swap prices and their data is within 1 bps with a standard deviation of 3 bps. Our estimates of the daily bid–ask spreads confirm that the inflation swap market is fairly liquid. We find that the United States and the United Kingdom's bid–ask spreads are of similar magnitude (1 to 2 bps, respectively) while the euro area and Japan have slightly larger bid–ask spreads (2 to 3 bps, respectively). Turning to the inflation-linked bond market, the United States is the largest inflation-linked government bond market globally with a market value outstanding of 871 billion dollars, followed by the United Kingdom with 559 billion dollars outstanding as of May 2012. In the euro area, France, Italy, and Germany are the largest issuers with 427 billion dollars combined outstanding as of May 2012. France's OATi are at 228 billion dollars, and Italy's BTPi are at 126 billion dollars, Germany' s Bundei/OBLei has 58 billion dollars outstanding followed by Japan JGBi with 51 billion dollars, see Norges Bank (2012) for additional details. The final institutional differences in our sample is taxation. In the euro area, Germany, France, and Italy apply same taxes as for the nominal bonds. UK real bonds are tax-exempted while in Japan the nonresidents are exempted from the tax while the domestic investors are subject to same tax as for the nominal bonds. In the United States, inflation-protected and nominal bond income are both taxable at the federal level.

4. RESULTS

Table 1 reports summary statistics for the pricing anomaly for each of the 47 sample pairs of inflation-linked and nominal bonds. The pricing anomaly reported in Table 1 is substantial. By country the Italian pairs exhibit the highest average mispricing of $8.71. The corresponding figures for Germany, United Kingdom, Japan, the United States, and France are $3.12, $1.93, $1.74, $1.67, and $1, respectively. The median values tend to be smaller, suggesting that the mispricing is right skewed, that is, there are larger positive mispricings than negative mispricings. Indeed, this appears to be the case as the maximum values are substantially greater than the (absolute value) of the minimum values, even though the average mispricing is relatively small. The average dollar mispricing for the United States is lower than the figure of \$3.13 reported by Fleckenstein, Longstaff, and Lusting (2014) for an earlier period. On average, nominal bonds are dearer than their synthetic counterparts. Among the 47 pairs, however, there are eight cases where the average daily price of the synthetic bond exceeds the average daily price of the nominal bonds. There are only four pairs for which the price of the synthetic bond never exceeded the price of the nominal bond. The standard deviation of the pricing anomaly tends to be relatively large suggesting that volatility in the mispricing might deter investors from engaging in this type of arbitrage strategy. This evidence motivates our investigation on the determinants of the pricing anomaly and the limits to arbitrage.[8]

To further examine the time-series properties of the average mispricing, Figure 1 plots the time series of the equally weighted-average dollar mispricing for all inflation-indexed and nominal bond pairs for each country. Figure 1 suggests that the mispricing is persistent, and is not a phenomenon associated solely with the financial crisis of 2008–09. While the mispricing peaks in the United States and Japan soon after the Lehman Default, we see no corresponding spike in the European countries. Nevertheless the peak of the mispricing appears to coincide with the Lehman Brothers default in autumn 2008. Table 2 reports the cross-correlations and autocorrelations for these aggregated series. In general the mispricings are correlated across the countries, with a typical correlation coefficient of around 0.5 or 0.6. There are some exceptions to this general observation though. The mispricing in Japan is negatively correlated with all other countries except Germany. Germany and France tend to have lower correlations with the other countries in the sample, around 0.2 on average. The lower panel of Table 2 reports the autocorrelations within each country. These autocorrelations are positive and greater than 0.6 at one lag for all countries. The series appears highly persistent. Autocorrelations are positive for over 12 months in the United States and United Kingdom, 6 months in Japan and France, and 4 months in Germany and Italy. One possible implication of this persistence is that the mispricing is not a short-term arbitrage opportunity that can easily be exploited but in fact is driven by persistent systematic risk factors. Furthermore, given the relatively low cross-country correlations of the mispricing, finding consistent systematic factors that correlate with the mispricing may shed light on whether this is an arbitrage opportunity or a compensation for risk.

These findings provide initial insights on the potential explanations for this pricing anomaly. To first determine whether the observed pricing anomaly correlates with risks in financial markets we run the following regression:

$$\Delta ln(mispricing)_{it} = \alpha + \beta x_{k,t} + \gamma_k + \delta_j + \varepsilon_{i,t}. \tag{1}$$

8. Table A2 in the Appendix reports more detailed information on the average mispricing for each pair examined in this study.

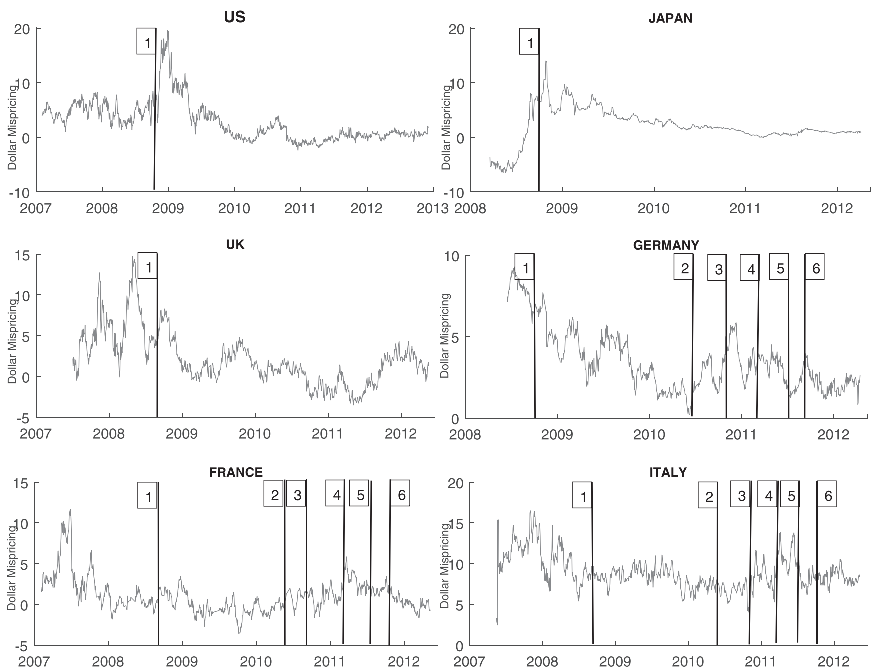

Fig 1. Time Series of Mispricing by Country.

NOTE: This figure plots the time series of the dollar nominal bond minus synthetic bond mispricing for all six countries in the study. From the top-left to the bottom-right are the United States, Japan, United Kingdom, Germany, France, and Italy. The mispricing is expressed in units of dollars per \$100 notional across the pairs included in the sample. The numbers in the boxes indicate important financial event dates. Box 1 is the Lehman Brothers default. Box 2 is the EU' s and IMF' s May 2010 agreement on a euro 110 billion bailout package for Greece. This bailout package was quickly replicated in support of Republic of Ireland in November of the same year, indicated by box 3 in the figure. Box 4 reports the decision of eurozone' s finance ministers adopted in February 2011 to set out a permanent bailout fund worth euro 500 billion, called the European Stability Mechanism (ESM). Box 5 reports the second bailout package for Greece worth euro 109 billion designed to resolve the Greek crisis and prevent contagion among other European economies. Finally, box 6 reports the Italy' s austerity budget of euro 50 billion adopted in September 2011. The sample period spans from February 2, 2007, to November 30, 2012.

The left-hand side variable is the change in the log mispricing variable defined as the log nominal bond price minus the log synthetic bond price. The right-hand side variables $x_{k,t}$ include the swap spread, the *VIX*, the 5-year *IRP*, the return on the global hedge fund index, and the bid-ask spread for the inflation-protected securities, γ_k is a country fixed effect and k indexes a specific country, δ_j is a year fixed effect and j indexes a specific year. Finally, i indexes a specific bond–synthetic bond pair and t represents a specific month.

Before we discuss the results of this regression, we believe it would be helpful to the reader to discuss a few aspects of our empirical strategy. First note that we run a pooled regression where we pool all bonds and countries together and try to explain the mispricing with risk factors that vary by country and month. We run our regression

TABLE 2

Aggregate Cross-Correlations and Autocorrelations of Mispricing

	US	UK	JPN	GER	FRA	ITA
United States	1.00	0.66	–0.32	0.19	0.24	0.54
United Kingdom	0.66	1.00	–0.21	0.06	0.06	0.50
Japan	–0.32	–0.21	1.00	0.39	–0.33	–0.07
Germany	0.19	0.06	0.39	1.00	–0.12	0.44
France	0.24	0.06	–0.33	–0.12	1.00	0.37
Italy	0.54	0.50	–0.07	0.44	0.37	1.00
	US	UK	JPN	GER	FRA	ITA
lag 1	0.87	0.83	0.94	0.73	0.77	0.62
lag 2	0.79	0.63	0.90	0.44	0.57	0.30
lag 3	0.72	0.54	0.92	0.40	0.50	0.19
lag 4	0.67	0.55	0.90	0.29	0.38	0.08
lag 5	0.59	0.49	0.84	0.12	0.23	0.06
lag 6	0.57	0.36	0.88	–0.05	0.22	–0.09
lag 7	0.54	0.25	0.82	–0.26	0.10	–0.15
lag 8	0.55	0.25	0.71	–0.34	–0.02	–0.10
lag 9	0.51	0.35	0.28	–0.25	0.02	–0.14
lag 10	0.55	0.35	–0.06	–0.19	–0.01	–0.20
lag 11	0.55	0.27	–0.26	–0.15	–0.12	–0.08
lag 12	0.60	0.17	–0.36	–0.10	–0.08	–0.04

Note: This table reports additional summary statistics for the mispricing. The first panel presents correlations across the countries; the second panel presents the autocorrelations within country. All data are averaged at the monthly level before calculating the correlations. The sample period spans from February 2, 2007, to November 30, 2012.

at the bond level so we can have more statistical power and because we believe that bond level is the correct unit of analysis as arbitrageurs will invest in specific bond pairs.[9] The second point we would like to make is that we do not observe risk factors directly, specifically the factors that are priced in financial markets. As a result, we must use common proxies for these risks and assume that there is a linear relationship between these proxies and risk. This approach has two limitations. The first is that we may omit relevant risk factors and so finding low or no correlation with the risk factors does not necessarily rule out that risk can explain the persistence of the mispricing. The other limitation is that we can discuss the extent to which our loadings on risk factors are statistically significant but we cannot directly interpret the magnitude of the coefficients.

We now turn our attention to the results presented in Table 3. We find that many of our proxies for risk factors are significantly correlated with the mispricing. We start with swap spreads, which have been long used as a measure of systemic credit and illiquidity risk on the financial system (see Duffie and Singleton 1997).[10] The

9. For readers concerned about the potential for serial correlation in our error terms, we conduct the test suggested by Wooldridge (2002, section 7.8.5 p. 177), a t-test on the AR(1) coefficient for the regression residuals. For our regression without country and year fixed effects we find a coefficient of -0.033, which is not significantly different than zero (p-value 0.54). For our regression with fixed effects we find a coefficient of -0.039 which is again not significantly different than zero (p-value 0.486). We thank an anonymous referee for bringing this point to our attention.

10. Other measures of systemic risk such as the spread between 3-month Libor rates and the overnight index swap (OIS) rate, the CDX index that captures the average CDS spread for investment grade bonds

TABLE 3

MISPRICING AND RISK FACTORS

	(1)	(2)	(3)	(4)
10-year swap spread	–0.001*	–0.012***	–0.001	–0.018***
	(0.001)	(0.004)	(0.001)	(0.005)
VIX	0.009***	0.008***	0.017***	0.016***
	(0.002)	(0.002)	(0.003)	(0.003)
Inflation risk premium	0.002**	0.008***	0.004	0.014***
	(0.001)	(0.002)	(0.002)	(0.004)
Hedge fund returns	–0.028	0.032	0.001	0.037
	(0.018)	(0.021)	(0.026)	(0.030)
Illiquidity	0.005	0.004	0.003	0.006
	(0.006)	(0.008)	(0.006)	(0.008)
Country FE		Yes		Yes
Year FE		Yes		Yes
Adj. R-squared	0.02	0.05	0.04	0.07
N	1,742	1,742	886	886
F stat	7.79	3.94	6.28	3.41
p-value F test	2.91e-07	1.07e-06	9.70e-06	0.0000367

NOTE: This table regresses the change in log mispricing (log nominal bond price minus log synthetic bond price) on a variety of explanatory variables. The explanatory variables are the *10-year swap spread* (10-year Swap Spread) for each country, *VIX* is the index of implied volatilities on equity index options for each country our proxy for the market's uncertainty, *Inflation Risk Premium* (*IRP*) is the 5-year *IRP* for each country and is estimated as the difference between the inflation swap and the expected inflation, as discussed in Section 3; the expected inflation for 5 years comes from University of Michigan survey for the United States; United Kingdom's expected inflation survey is reported by BoE; Japan used NATAKA survey reported in Bank of Japan; and finally for the EU countries the expected inflation is the ECB survey. *Hedge Fund Returns* is the HFRX Hedge Fund global index return. Finally, *Illiqudity* is the bid–ask spreads of the inflation-indexed bonds of each security in our sample. Country FE denotes if country fixed effects are used to account for country specific factors that are constant over time. Year FE denotes if year fixed effects are used to account for time-specific factors that are constant across countries. F stat is the F-statistic for the null hypothesis that all the coefficient are zero. p-Value F test is the p-value for this test. Column (1) is our baseline regression. Column (2) adds year and country fixed effects. Columns (3) and (4) repeat the analysis of 2 and 3 without the U.S. observations. Significance levels: * : 10%, ** : 5%, and *** : 1%.

swap spread enters negatively. We view an increasing swap spread as indicating reduced demand for corporate securities and increased demand for sovereign securities. This demand flows asymmetrically into inflation-protected securities, naturally lowering the mispricing. Second, the *VIX* enters positively. We again interpret this as arbitrageurs being exposed to risk, in this case volatility risk, which increases the mispricing when the risk rises. Similarly, the *IRP* enters positively. When investors are willing to pay more to insure against inflation risk the mispricing widens. This result suggests that the arbitrage strategy is exposed to short-term inflation risk. Intuitively this makes sense, in regimes of increased uncertainty investors are willing to pay more to insure against inflation risk. In these states the mispricing widens.

On the other hand, we find no significant evidence that hedge fund returns correlate with the mispricing. We will explore this proposition in more detail in the paper and again we will find little support for the slow-moving capital hypothesis to explain the mispricing. Furthermore, we do not find significant evidence that the reported pricing differential correlates to liquidity risk in the market for inflation protection as proxied by linkers' bid–ask spreads. This evidence together with the large standard deviations of the mispricing reported in Table 1 cast doubt on the view that this is an

result highly correlated with the swap spreads and do not provide a significant incremental contribution in explaining the relation of the pricing anomaly with the macrofinancial systemic risk.

easily exploitable arbitrage opportunity limited only by arbitrageurs lack of access to capital. In column two of Table 3 we present the same results controlling for country and time (year) fixed effects and the results are very similar. The fixed effects are able to control for country-specific factors that are constant over time. This would include, for example, institutional factors that are specific to the countries we examine.

Since the majority of our bonds are U.S. bonds one might worry that these observations drive our results. To address this potential concern, we examine whether our results are sensitive to excluding the observations on U.S. bonds. Columns 3 and 4 of Table 3 repeat the previous analysis without the United States. The results do not change appreciably. When looking at our fixed effect analysis we find that the coefficients on the swap spread, *VIX* and *IRP* are all still significant and in fact larger in magnitude.

One potential issue with our results is the low R-squared values from our models. We do not find this surprising because our outcome variable is at the bond level and our explanatory variables are at the country level. Indeed, when we later aggregate the data by month (in Section 4.2) and rerun the regression, we find that the R-squared values are an order of magnitude larger. However, it is useful to check that our regression model does have statistically significant explanatory power for the mispricing. Table 3 reports the F statistic for the test that all the coefficients on the explanatory variables equal zero along with the corresponding p-values. The largest p-value is 0.00004 meaning we can reject the hypothesis that our model does not explain the mispricing with an extremely high degree of statistical confidence.

Next, we explore the role of country default risk in explaining the mispricing. We conjecture that if the pricing differential accounts for a premium, in case the issuer fails to meet her obligations, then this effect should be reflected in its correlation with the country-specific CDS premium. In Table 4, column 2, we rerun our baseline regression using the level of the mispricing but now subtract off the CDS premium for insuring against sovereign default from the mispricing. We do this to see if any of the above identified risk factors are proxying for exposure to default risk. CDS spreads should capture all relevant information concerning the altered risk of default for each country. In addition, CDS spreads should also capture the impact of the adopted policy measures such as the ECB' s securities market programme (SMP) or any rescue loan supplied to financially distressed countries on the bond markets. The results are in Table 4. One can see that the coefficients from the two regressions are very similar, which suggests that default is not an important determinate of the mispricing. However, while the *VIX* variable was significant in the previous regression, they are no longer significant once we control for sovereign default risk through the CDS premium. This result suggests that volatility is correlated with sovereign default risk. We also find that the swap coefficient is smaller and significantly different than in the regression not factoring in CDS premium. Part of the mispricing premium appears due to sovereign default risk that lessens in the presence of stronger forecasted economic activity.

However, quantitatively the CDS premium is small relative to the mispricing. Figure 2 plots the monthly average of the mispricing and the monthly average of

TABLE 4
Mispricing and Default Risk

	(1)	(2)
10-*year Swap Spread*	–1.306***	–0.973***
	(0.221)	(0.209)
VIX	0.733***	0.035
	(0.237)	(0.223)
Inflation risk premium	0.781***	0.722***
	(0.151)	(0.142)
Hedge fund returns	3.041	1.743
	(2.241)	(2.113)
Illiquidity	0.028	0.342
	(0.564)	(0.532)
Constant	3.591***	2.719***
	(0.586)	(0.553)
County FE	Yes	Yes
Year FE	Yes	Yes
Adj. R-squared	0.05	0.03
N	1,742	1,742

Note: This table regresses the mispricing, nominal bond price minus synthetic bond price, column (1), and the difference between the pricing differential and the CDS value for insuring against sovereign default, column (2), on the explanatory variables in Table 3. Significance levels: * : 10%, ** : 5%, and *** : 1%.

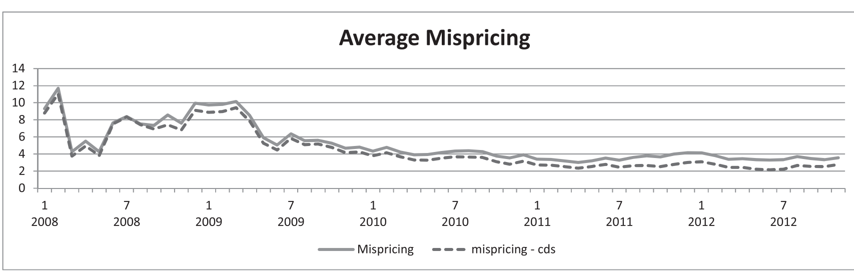

Fig 2. Mispricing with and without CDS.

Note: This figure plots the time series of the average dollar mispricing (solid line) for all six countries in the study in the top panel. The dashed line is the difference between the pricing differential and the CDS spreads for all six G7 countries. The sample period spans from February 2, 2007, to November 30, 2012.

the mispricing minus the CDS variable. One can see that the plots are almost identical whether or not the CDS premium is subtracted from the mispricing or not. This result suggests strongly that default risk—even in Europe where default was seen as a real possibility—is not the reason that there is mispricing between inflation indexed and nominal government bonds. If investors were concerned about default risk they could purchase CDS insurance for their portfolio and still make almost the same arbitrage profit.

We also conjecture that macro-economic risks factors will correlate with this pricing anomaly. Specifically, in periods of increased *inflation expectations* the demand for the relatively cheap inflation-protected securities will rise narrowing the pricing anomaly. On the other hand, in periods of expected deflation the demand will

TABLE 5

Macrovariables and Inflation Expectations

	(1)	(2)
10-*year swap spread*	–0.012***	–0.011**
	(0.003)	(0.004)
Hedge fund returns	0.022	0.0439*
	(0.020)	(0.023)
VIX	0.005*	0.003
	(0.003)	(0.003)
Inflation risk premium	0.007***	0.009***
	(0.002)	(0.002)
Illiquidity	0.007	0.009
	(0.008)	(0.008)
Δ*log(Oil price)*	–0.021**	–0.0221**
	(0.009)	(0.010)
Δ*Overnight bank lending rate*	0.015***	0.010*
	(0.005)	(0.006)
Δ*log(Industrial production)*	0.021	0.061
	(0.038)	(0.058)
Δ*Government budget deficit*	–0.001	–0.002
	(0.001)	(0.001)
Δ*Median inflation expectations*		–0.004
		(0.005)
Δ*Inflation uncertainty*		0.001
		(0.001)
Δ*Inflation disagreement*		0.002
		(0.003)
Constant	0.035***	0.030**
	(0.010)	(0.012)
Country FE	Yes	Yes
Year FE	Yes	Yes
Adj. R-squared	0.06	0.07
N	1,742	1,428

Note: This table replicates the results in Table 3 adding macro-economic variables to examine the exposure of the mispricing to macro-economic risk factors. The macro-economic variables in this table are: Δlog(*Oil Price*) denotes log-changes in crude oil (West Texas Intermediate) spot prices. Δ*Overnight Bank Lending Rate* is the *overnight bank lending rates* which for the Fed Funds are rates from FRED for the United States, for Japan, it is the basic loan rate, for the United Kingdom and the EURO area countries we use Libor and Euribor. Δlog(*Industrial Production*) is the *industrial production*. *Government deficits* (Δ*Government Budget Deficit*) are used as they can be associated with larger default risk and affect bond prices. *Inflation expectations* (Δ*Median Inflation Expectations*) are used here because increased *inflation expectations* could increase demand for inflation-protected securities and also be consistent with an improved outlook on the economy. Finally, Δ*Inflation Uncertainty* is defined as the mean of the standard deviation of each forecaster's forecast, $E[\sigma_i(\pi)]$ for next year and Δ*Inflation Disagreement* is defined as the standard deviation of each forecasters expected inflation $\sigma(E_i\pi)$. Δ denotes the first difference of the variables. Column (1) adds macro-economic variables to the baseline regression, column (2) adds inflation expectation variables. Significance levels: * : 10%, ** : 5%, and *** : 1%.

switch leading to a widening of the mispricing. In Table 5 we augment our baseline regression with several macro-economic variables: *oil prices*, short-term interest rates, *government deficits*, and survey-based *inflation expectations*. We find that increased *oil prices* are correlated with a reduction in the mispricing. We interpret this result as higher *oil prices* being associated with an increase in world demand. This increase leads investors to expect stronger economic activity going forward reducing the risk exposure of the mispricing strategy. While it is possible that increased *oil prices* might lead to increased *inflation expectations*, and that might affect the mispricing, we find below that changes in *inflation expectations* do not effect the mispricing. Therefore, we interpret changes in *oil prices* as capturing the state of the economy. Other macro-economic variables are not correlated with the mispricing.

TABLE 6

Dynamic Relationship between Factors and Mispricing

	(1)	(2)
10-*year swap spread*$_{t-1}$	–0.000	–0.002
	(0.001)	(0.004)
VIX$_{t-1}$	0.006**	0.005**
	(0.002)	(0.002)
Inflation risk premium$_{t-1}$	0.002*	0.004**
	(0.001)	(0.002)
Hedge fund returns$_{t-1}$	–0.013	0.020
	(0.021)	(0.020)
Illiquidity$_{t-1}$	0.009	0.014*
	(0.006)	(0.007)
Constant	–0.001	0.011
	(0.002)	(0.012)
Country FE	No	Yes
Year FE	No	Yes
Adj. *R*-squared	0.01	0.02
N	1,742	1,742

Note: This table examines the dynamic relationship between the mispricing and the lag of relevant risk factors. The explanatory variables are the same as in Table 3. The time subscripts $_{t-1}$ denotes lag one in variables. Column (1) and (2) both use the log change in mispricing as the dependent variable. Column (2) adds country and year fixed effects. Significance levels: * : 10%, ** : 5%, and *** : 1%.

Deficits, short-term interest rates, and *industrial production* all enter the regression insignificantly as do 1-year-ahead median *inflation expectations* and the measures of disagreement and uncertainty.[11] It appears that with the exception of *oil prices*—financial market variables as opposed to more general macro-economic variables are important in determining the mispricing.

Finally in Table 6, we look at the dynamic relationship between our risk factors and the mispricing. Specifically we regress the mispricing on the lag values of the factors. We find that increased volatility and lower liquidity lead to an increase in the mispricing. This result is consistent with increased volatility leading investors to reduce their exposure to the mispricing strategy. Additionally, decreased liquidity (through higher bid-ask spreads) leads to a lower return from the arbitrage strategy. This effect would lead to fewer investors exploiting the arbitrage and a widening of the mispricing. Again the *IRP* is positive suggesting that the mispricing is exposed to increased inflation risk leading the mispricing to widen.

In Table 7 we regress the factors we used to explain the mispricing on the nominal bond and the synthetic bond separately. If there were no arbitrage opportunity then the factors should have an equal effect on both the nominal and the synthetic bond. First, we see that an increase in the swap spread leads to increased prices in the synthetic bond market but has little to no effect in the nominal bond market. Similarly the *VIX* has a large positive effect on nominal bond prices but no effect on the synthetic bond

11. *Inflation expectations* are based on survey data. Each forecaster reports a probability distribution over possible future values of inflation. From these data we calculate the forecaster's expected inflation $E_i\pi$ and the standard deviation of the agents forecast $\sigma_i(\pi)$. Disagreement is defined as the standard deviation of each forecasters expected inflation $\sigma(E_i\pi)$ and uncertainty is defined as the mean of the standard deviation of each forecaster's forecast, $E[\sigma_i(\pi)]$.

TABLE 7

Factors on Nominal and Synthetic Bond

	(1) Nominal	(2) Synthetic	(3) Mispricing
10-year swap spread	–0.004	0.008**	–0.012***
	(0.003)	(0.004)	(0.002)
VIX	0.007**	–0.001	0.008***
	(0.003)	(0.004)	(0.002)
Inflation risk premium	0.001	–0.007***	0.008***
	(0.002)	(0.002)	(0.001)
Hedge fund returns	0.001	–0.031	0.032
	(0.028)	(0.036)	(0.021)
Illiquidity	–0.021***	–0.025***	0.004
	(0.007)	(0.009)	(0.005)
Constant	0.016**	–0.016*	0.032***
	(0.007)	(0.009)	(0.005)
Country FE	Yes	Yes	Yes
Year FE	Yes	Yes	Yes
Adj. R-squared	0.02	0.01	0.05
N	1,742	1,742	1,742

Note: This table decomposes the effect on the mispricing into effects on the nominal bond and the synthetic bond. The explanatory variables are same as in Table 3. Column (1) uses the log change in the nominal bond price as the dependent variable, column (2) uses the log change in the synthetic bond price as the dependent variable and column (3) uses the difference in these log prices as the dependent variable. Significance levels: * : 10%, ** : 5%, and *** : 1%.

market. This may be a flight to safety effect that widens the mispricing consistent with Longstaff (2004), Krishnamurthy (2002), and Bansal et al. (2010), who argue that investors value the liquidity and safety of treasury bonds, that is, the liquidity preference theory.

To investigate whether policy actions such as changes in the short-term interest rates to maintain inflation targets affect the pricing differential, Figure 3 plots the response of the mispricing to an increase in the short-term interest rate. If the pricing differential is a result of a pure arbitrage opportunity, we expect changes in the policy measures to have a marginal impact on the mispricing or perhaps widen the mispricing as the cost of funds to arbitrageurs increases. However, if the pricing differential acts as compensation for bearing inflation risk, changes in policy actions should have an effect on them. An unanticipated increase in interest rates may signal that policymakers expect inflation to be higher and the economy to be stronger in the future. As a result, the riskiness of the arbitrage strategy has been reduced.

We identify this change from a structural VAR, the estimation procedure of which we describe in Appendix B. We find that the increase in the short-term rate lowers the mispricing in the short run (1–5 months) but in the long run (1 year) leads to an increase in the mispricing. Presumably, an increase in the short-term interest rate affects the nominal bond market more than the synthetic bond market leading to a larger fall in nominal bond prices. However, in the long run, prices rebound and the mispricing ends up higher than before the shock. A possible interpretation of these results is that in the short run, an unexpected increase in the short-term interest rate raises *inflation expectations*, which lowers the mispricing through increased demand for inflation-protected securities. However, eventually the increased interest rates lower economic

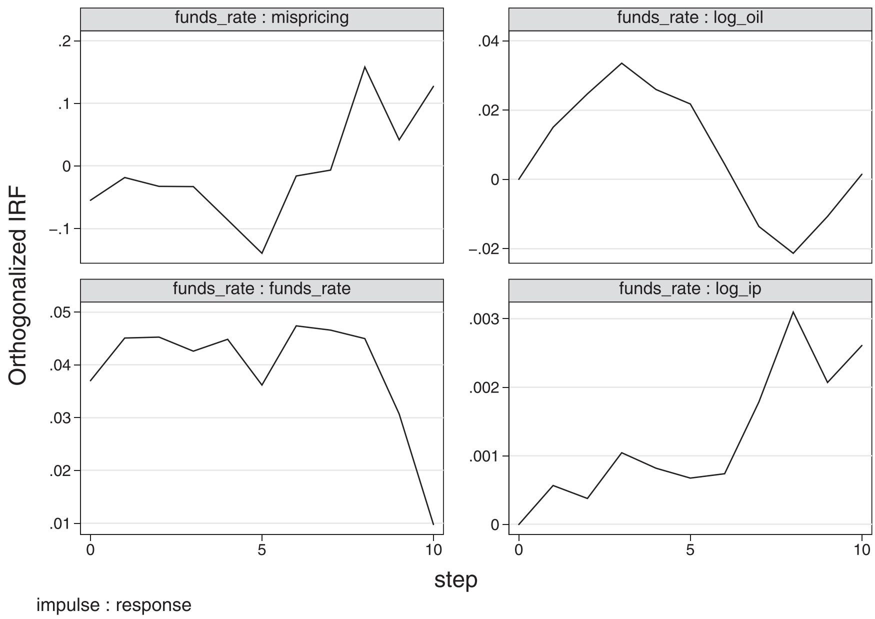

Fig 3. Response of the Mispricing to an Increase in Short-Term Interest Rates.

NOTE: This figure plots the response of the mispricing to an increase in the short-term interest rate. We identify this change from a structural VAR the estimation procedure of which we describe in Appendix B. The sample period spans from February 2, 2007, to November 30, 2012.

activity and inflation, as evidenced by lower *oil prices*, leading to a rebound in the size of the mispricing. These results are consistent with our original conjecture concerning the effect of macro-economic variables on the mispricing. Increased *inflation expectations* lower the mispricing as investors demand inflation protection to hedge against an inflation rise, however, when *inflation expectations* subside with weakened economic activity, the demand dissipates leading to an increase in the mispricing.

4.1 Slow-Moving Capital and Institutional Explanations

One proposed explanation for the limits of arbitrage is the lack of capital to narrow the arbitrage opportunity to zero. According to this slow-moving capital theory, when more capital becomes available to arbitrageurs we should see a narrowing in the mispricing. Table 8 regresses the change in mispricing on lag returns (four) of various hedge fund indices. The indices represent global, macrostrategy, relative value, convertible arbitrage, volatility, high yield, and fixed income sovereign hedge fund returns. We find no consistent evidence that past positive hedge fund returns result in lower mispricing. Of the six significant returns, four are negative and two are positive. Importantly, we find no evidence that sovereign or relative value hedge funds returns

TABLE 8
HEDGE FUND REGRESSIONS

	(1) Global	(2) MCR	(3) RV	(4) Conv	(5) Vol	(6) Yield	(7) Sov
lag 1	–0.032	–0.003	–0.040	–0.057	0.026	–0.035*	–0.078
	(0.026)	(0.031)	(0.045)	(0.043)	(0.022)	(0.021)	(0.056)
lag 2	–0.046**	–0.014	–0.024	–0.001	–0.127***	–0.044	–0.059
	(0.022)	(0.036)	(0.034)	(0.028)	(0.030)	(0.031)	(0.040)
lag 3	0.056**	0.017	0.016	0.009	0.044**	–0.008	0.007
	(0.024)	(0.026)	(0.041)	(0.035)	(0.021)	(0.020)	(0.034)
lag 4	–0.000	–0.073**	0.023	0.019	0.060***	0.019	0.005
	(0.017)	(0.032)	(0.027)	(0.019)	(0.019)	(0.022)	(0.026)
Constant	–0.000	–0.000	–0.000	–0.000	–0.000	–0.000	–0.000
	(0.000)	(0.001)	(0.000)	(0.000)	(0.000)	(0.001)	(0.000)
FE	No	No	No	No	No	No	No
Adj. R^2	0.01	0.01	0.01	0.02	0.02	0.01	0.02

NOTE: This table regresses future mispricing on lag changes in various hedge fund returns. All hedge fund returns are subcategories of the HFRX Index. For dependent variables, column (1) uses the HFRX global index returns, column (2) uses the HFRX Macro-Strategy Index, column (3) uses the HRFX Relative Value Arbitrage Index column (4) uses the HFRX Fixed-Income Convertible Arbitrage Index, column (5) uses the HFRX Volatility Strategies Index, column (6) uses the HFRX Fixed Income Alternative Yield Index and finally, column (7) uses the HFRX Fixed-Income Sovereign Index. Significance levels: * : 10%, ** : 5%, and *** : 1%.

are correlated with the mispricing. These results cast doubt on slow-moving capital to explain the mispricing.

4.2 Comparison with Fleckstein et al. (2014)

As discussed earlier, we follow closely the seminal paper of Fleckenstein, Longstaff, and Lusting (2014) [FLL] in constructing the arbitrage strategy. FLL documents the arbitrage opportunity and shows that it holds even after accounting for common explanations like taxes, trading costs, market microstructure considerations, and liquidity of inflation linked bond. They also show, via a simple regression analysis, that swap spreads are not significantly related to the mispricing but hedge fund assets and repo fails (a measure of liquidity) are. These are different conclusions from the results reported in Table 3 and therefore warrant further investigation.

Differences in our results and the ones reported in FLL are to be expected given the differences in sample size and time period as well as the different approaches adopted in these papers, therefore we do not view them as surprising. To list some of the differences between these papers: (i) their analysis is conducted at the monthly aggregate level while ours is conducted at the monthly bond level, (ii) we have data on six countries while their analysis is only for the United States, (iii) their time period ends in 2009 while ours continues to 2012, (iv) we use the level of the swap spread while they use the change, (v) they use global hedge fund assets while we use returns, and (vi) they use repo fails as a liquidity measure while we use bid ask spreads. Although it is not possible to replicate their analysis entirely, for example, we cannot get repo fail data by country, we can try to move our analysis closer to theirs and see if and how our results change.

TABLE 9

Comparison with Fleckstein Et Al. (2014)

	(1) baseline	(2) time period	(3) bp	(4) bp (FE)	(5) Δswap	(6) US	(7) monthly
10-*year spread*	–0.001*	–0.006*	–0.139*	–1.306***	–0.011***	–0.007	–0.029*
	(0.001)	(0.003)	(0.077)	(0.401)	(0.004)	(0.004)	(0.015)
VIX	0.009***	0.002	0.900***	0.733***	0.006***	0.002	–0.021
	(0.002)	(0.010)	(0.220)	(0.237)	(0.002)	(0.002)	(0.015)
Inf. risk prem.	0.002**	0.005*	0.243*	0.781***	0.002	0.004*	–0.020
	(0.001)	(0.003)	(0.138)	(0.213)	(0.001)	(0.002)	(0.013)
Hedge Ret.	–0.028	–0.160	–3.332*	3.041	–0.019	–0.039*	–0.105
	(0.018)	(0.097)	(1.974)	(2.361)	(0.019)	(0.020)	(0.194)
Illiquidity	0.005	0.074	0.300	0.028	0.005	–0.022	–0.725**
	(0.006)	(0.076)	(0.429)	(0.550)	(0.006)	(0.043)	(0.294)
Country FE				Yes			
Year FE				Yes			Yes
Constant	0.002	0.010	0.234	3.591***	–0.001	0.018	0.192***
	(0.002)	(0.008)	(0.171)	(1.180)	(0.001)	(0.013)	(0.036)
Adj. R^2	0.02	0.05	0.02	0.05	0.03	0.03	0.79
N	1,742	327	1,742	1,742	1,742	856	59

Note: This table redoes our baseline analysis of Table 3 but modifying the sample to be more inline with Fleckstein et. al (2014) [FLL] Table V. Column (1), baseline, is the regression from Table 3. Column (2), time period, uses data only through 2009 consistent with the sample in FLL. Column (3), bp, uses the first difference of the dollar value of the mispricing per $100 dollars notional consistent with FLL. Column (4) adds fixed effects to this the regression of column (3). Column (5), Δ*swap*, uses the first difference of the 10 year swap spread consistent with FLL. Column (6), US, keeps only observations on U.S, bonds as in FLL. Column (7), monthly, aggregates the U.S. sample by month consistent with FLL. Significance levels: * : 10%, ** : 5%, and *** : 1%.

We conduct this analysis in Table 9. We compare the baseline regression of Table 3 to various alternative analyses, which move us closer to FLL. To this end, we compare five alternatives. First, we restrict our sample to stop at November 2009, consistent with FLL. We see that in this time period our *VIX* coefficient is no longer significant suggesting that the time period may partially explain the inability of FLL to find risk factors' importance in explaining the mispricing. Second, we express the mispricing variable as a change in the dollar value, consistent with FLL. The only difference we find relative to our baseline result is that the hedge fund return coefficient is significant and negative. However, once we include year fixed effects that effect disappears. We interpret this as a cautionary tale when one relies solely on a time-series analysis. It is quite possible that hedge funds do well when risk factors subside and that is also the time when the mispricing narrows. However, as the fixed effect controls demonstrate, one cannot interpret this as a causal effect. Third, we examine putting in the swap spread variable in changes versus in levels. We find that the swap spread is still significant. Fourth, we restrict the sample only to the United States. The swap spread is no longer significant nor is the *VIX* effect. This analysis suggests that the risk factors are more prominent outside the United States and restricting the analysis only to the United States limits the analysis. Finally, in the fifth control we run our regressions at an aggregate monthly level and we include only the U.S. pairs up to November 2009, the closest to FLL. The swap spread is again significant, albeit the other variables are not, while the illiquidity effect changes sign. This final result suggests that aggregating the mispricing may understate the risk of the arbitrage strategy especially if these risk factors effect the tails of the mispricing distribution in a way that is not apparent in the middle of the distribution.

5. EUROZONE CRISIS

We view the eurozone crisis as an ideal environment to study the effects of the risk factors on the pricing differential. The euro area is informative due to the existence of numerous competing sovereign issuers—with different credit ratings and associated default probabilities—that issue obligations in the same currency, therefore the impact on yields of a fall in the credit rating of a particular issuer can be marked. The time period that our sample covers also lends well for this analysis as it covers the pre- and- post–general financial markets distressed period and the euro crisis period including the late 2012 when the ECB's and other policy interventions appeared to have stabilized the credit market in the eurozone. Accordingly, we expect that the macrofinancial, macro-economic and policy measures to have substantially different effects on the eurozone pricing differential than with the non-eurozone pairs analyzed in this paper and to examine the behavior of the pricing anomaly in an environment with real economic tail risk and strong deflationary pressures.

The average pricing anomaly for the eurozone pairs is about $4, which is considerably higher than the $1.67 for the non-eurozone pairs. Figure 4 plots the time series of average and aggregate dollar mispricing for the eurozone countries and the average and aggregate mispricing for the non-eurozone countries. During 2011–12, when the crisis of confidence surrounding the euro was at its peak the average mispricing for the eurozone countries is substantially higher than the average for the non-eurozone countries. Take Italy for example, whose secondary government bond market has the largest outstanding amount in the eurozone.[12] There the average mispricing jumps from $7.8 for May 2008 to December 2010 to $10.74 for May 2011 to August 2012, and then drops to $3.31 for September to December 2012. This change of the mispricing for Italy coincides with rising sovereign credit risk in eurozone countries under financial stress and the CDS and bond market diverging signals as reported by Moody's on December 21, 2010, and February 24, 2011. Successively, Moody's on June 17, 2011, places Italy's Aa2 rating on review for possible downgrade and effectively downgrades it to A2 with negative outlook on September 16, 2011. A reversion for the eurozone countries during the latter stages of 2012 coincides with a strengthening of support for the euro on the part of the ECB.[13] The corresponding figures for the same time period for the other two eurozone countries are $3.23, $2.82, and $2.45 for Germany and $1.13, $1.79, and $0.89 for France. The average mispricing for all

12. Data on Italian bond market can be found at: https://www.mtsmarkets.com/data-and-participant-reports/market-data-reports

13. On March 05, 2012, the ECB provided additional 3-year funding for the eurozone and on July 30, 2012, the governor of ECB Mario Draghi reassured the markets that ECB will continue with the support, but also warned that ECB cannot resolve the debt crisis. See https://www.moodys.com/credit-ratings/Italy-Government-of-credit-rating-423690 for Italy's credit rating.

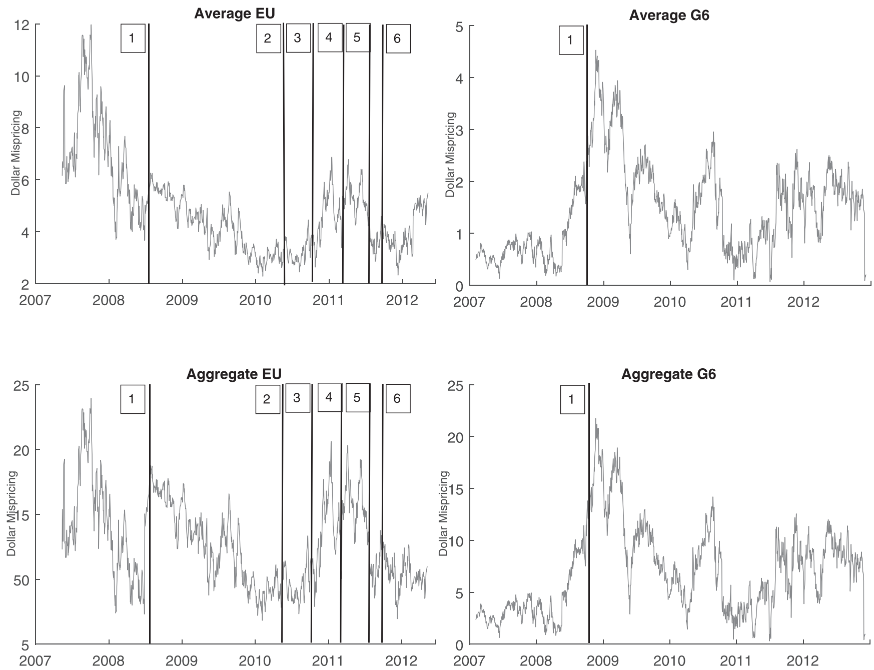

Fig 4. Mispricing by Eurozone and Non-Eurozone.

NOTE: This figure plots the average (top two figures) and aggregate (bottom two figures) mispricing for the eurozone and the G6 countries separately. The sample period spans from February 2, 2007, to November 30, 2012.

three eurozone countries for May 2008 to December 2010 is \$5.31, for November 2011 to August 2012 is \$4.27 and September 2012 to December 2012 is \$4.74. The corresponding figures for average mispricing for the non-eurozone countries were lower, and more stable, throughout this period, at \$1.88, \$1.79, and \$1.48, respectively.

To more formally investigate the change in the mispricing during the eurocrisis for the eurozone countries relative to the non-eurozone countries, we run a difference-in-difference regression

$$\begin{aligned} Mispricing_{i,t} &= \alpha + \beta * EuroCrisis_t + \gamma * EuroCountry_k \\ &\quad + \delta EuroCrisis_t * EuroCountry_k + \varepsilon_{i,t}. \end{aligned} \tag{2}$$

Here mispricing is expressed as dollars per \$100 notional value, *EuroCrisis* is an indicator variable that the date is May 2010 or later (the date of the EU/IMF bailout of Greece), *EuroCountry* is an indicator variable that the country uses the euro (i.e., Germany, France, or Italy), and *EuroCrisis*×*EuroCountry* is the product of these two variables.

TABLE 10

Euro-Crisis Difference-in-Difference Regression

	(1)	(2)
Euro crisis	–4.389***	
	(0.452)	
Euro country	0.643	
	(0.767)	
Euro crisis × *Euro country*	3.387***	3.217***
	(0.852)	(0.775)
Constant	4.777***	6.608***
	(0.424)	(0.836)
Country FE		Yes
Year FE		Yes
Adj. R-squared	0.11	0.23
N	1,790	1,790

Note: This table regresses, in column (1), the mispricing (here expressed as dollars per $100 notional value) on *Euro Crisis*, an indicator variable that the date is May 2010 or later, *Euro Country*, an indicator variable that the country uses the euro (i.e., Germany, France, or Italy), and the product of these two variables. Column (2) adds country and year fixed effects. Significance levels: * : 10%, ** : 5%, and *** : 1%.

To interpret this regression note that each coefficient captures an average of the mispricing: α is the non-euro mean before the crisis, $\alpha + \beta$ is the non-euro mean after the crisis, $\alpha + \gamma$ is the euro mean before the crisis, and $\alpha + \beta + \gamma + \delta$ is the euro mean during the crisis. Our coefficient of interest is δ. This captures the change in the mispricing in the euro area during the crisis relative to the non-euro area.

The results from the regression are in Table 10. We find that $\delta = 3.39$ (3.22 with country and year fixed effects).[14] This result is statistically significant and indicates that the mispricing increased by 3 dollars per 100 dollars notional relative to the non-euro area during the crisis. Additionally, note that the *EuroCrisis* coefficient is negative. *EuroCrisis* is a time dummy so this coefficient indicates that over time the mispricing has fallen in the sample. Finally, the *EuroCountry* dummy is statistically insignificant suggesting that the difference in the mispricing value between the euro and non-euro area was not important before the *EuroCrisis*.

During the eurozone crisis risk factors associated with the mispricing strategy: default risk, downside economic risk, deflation risk, were all more pronounced. If the mispricing between the nominal and synthetic bonds represents a compensation for risk then we would expect the mispricing to be larger and more sensitive to risk factors in the eurozone countries particularly during the eurozone crisis. This indeed seems to be true. Table 11 redoes the analysis in Table 3—which examined the factors that correlated with the mispricing—restricting the regression to only the eurozone countries: Italy, France, and Germany. When we restrict the regression to the eurozone countries, the signs and significance of the coefficients do not change. However, the magnitudes become larger. For instance, the coefficient on the 10-year swap spread is -0.015 versus -0.012 for all countries. More to the point, the coefficient on the *VIX* is 0.014 and the *IRP* is 0.023 versus a value of 0.008 for both the *VIX* and *IRP*

14. Note that *EuroCrisis* and *EuroCountry* drop out of the fixed effects regression because they are subsumed into the country and year fixed effects.

TABLE 11

MISPRICING IN THE EUROZONE COUNTRIES

	(1)	(2)
10-year swap spread	–0.014**	–0.015**
	(0.006)	(0.007)
VIX	0.013***	0.014***
	(0.004)	(0.004)
Inflation risk premium	0.024***	0.023***
	(0.005)	(0.006)
Hedge fund returns	0.055*	0.080**
	(0.033)	(0.038)
Illiquidity	0.009	0.012
	(0.006)	(0.009)
Country FE	No	Yes
Year FE	No	Yes
Constant	0.029**	0.031**
	(0.014)	(0.015)
Adj. -squared	0.10	0.11
	572	572

NOTE: This table replicates the results for Table 3 restricting the sample to only the eurozone countries (France, Germany, and Italy). The explanatory variables are same as in Table 3. Column (1) is our baseline regression; column (2) adds in year and country fixed effects. Significance levels: * : 10%, ** : 5%, and *** : 1%.

coefficients for all the countries. Again, our measure of liquidity, the bid–ask spread is not significant. The one clear difference between the eurozone regression and the baseline regression is that the hedge fund returns are now positive and significant. This suggests if anything hedge funds are exacerbating the mispricing as opposed to arbitraging it away.

An important question is if the difference in magnitude of the regression coefficients between the eurozone and non-eurozone is statistically significant. To formally test this question we have rerun our regression on the full sample with all the explanatory variables interacted with an indicator variable for being a euro-country. The t-test on the interaction coefficient allows us to test if the coefficients are different across the euro and non-eurozone. We have omitted the table due to space constraints but to convey the main results, we found that the difference in the coefficient on the swap spread, *IRP* and hedge fund returns are statistically different but the coefficient on the *VIX* is not. To summarize, the pricing anomaly is more pronounced in the eurozone area. This result is consistent with the mispricing being a premium for taking on the risk associated with the possibility of persistent weak economic activity resulting from the ongoing euro crisis and fiscal consolidation in the eurozone.

6. CONCLUSION

We report new evidence that the pricing differential between sovereign nominal bonds and synthetic bonds that replicate nominal bonds' cash flow is positive and persistent in all six of the countries that we analyzed. This mispricing occurs because the break-even inflation rate differs from the inflation rates implied by the swap market. We found that this mispricing correlates with volatility risk, inflation risk, and downside economic risk. We found little evidence that increasing the capital available to arbitrageurs reduced this mispricing. The mispricing was larger in the eurozone as was the magnitude of its correlation with the relevant risk factors. We interpret these results as being consistent with the mispricing being a compensation for risk.

Our results are robust to a large number of samples and specifications. We conduct our analysis for three groups: the full sample, the eurozone, and all countries except the United States. We also designed the analysis and robustness checks to ensure that country idiosyncrasies do not drive the results. For example, we used country fixed effects to control for unobservable country-specific features and year fixed effects to control for unobservable year effects. We also reran the results without the bond pairs with the highest maturity mismatch, without Germany and France who pay annual coupons, and without the United Kingdom and Japan who have no deflation floor. We found the in each case the results were consistent with our baseline regression.

A reader may note that we do not conclude that there are important financial limits to this arbitrage, and it may then seem puzzling as to why this mispricing persists. However, although the arbitrageur could lock in a profit if she holds the bonds to maturity, we have shown that the mispricing widens with an increase in the level of important risk factors. It therefore may be the riskiness of the strategy, which limits the degree to which arbitrageurs can arbitrage away the mispricing. We therefore view the mispricing as a compensation for bearing this risk. Our interpretation of this finding is therefore closer in spirit to Shleifer and Vishny's (1997) "limits to arbitrage" stemming from institutional features. In this view, investors may not be able to hold these securities to maturity. Over time the mispricing may widen and they may face margin calls forcing them to liquidate the position at a loss and at a time when risk factors are particularly high.

While the mispricing variable we estimate is clearly not a measure of *inflation expectations*, our paper is informative to the literature that uses nominal and inflation-protected bonds to measure the expected inflation rate. A natural starting point in the measurement of *inflation expectations* is the break-even inflation rate, the difference in yields, on matched nominal and inflation-protected bonds. However, as noted by many academics and practitioners these break-even inflation rates differ markedly from other measures of *inflation expectations*, particularly inflation swaps. One approach in the literature is to use structural models with an additional risk premium for holding inflation-protected securities to back out the true inflation expectation. Our results are generally supportive of this approach.

Moving forward, our paper is sympathetic to the general notion that asset prices can be an important way to measure expectations, not only for inflation but for measures of future asset prices, economic activity, and interest rates. It suggests that features like segmented markets are less important in determining the asset prices and that potential arbitrage opportunities are more likely compensations for taking on risk. Consequently, reliable information on expectations can be extracted from financial markets with careful economic and financial modeling.

APPENDIX A: ADDITIONAL TABLES & FIGURES

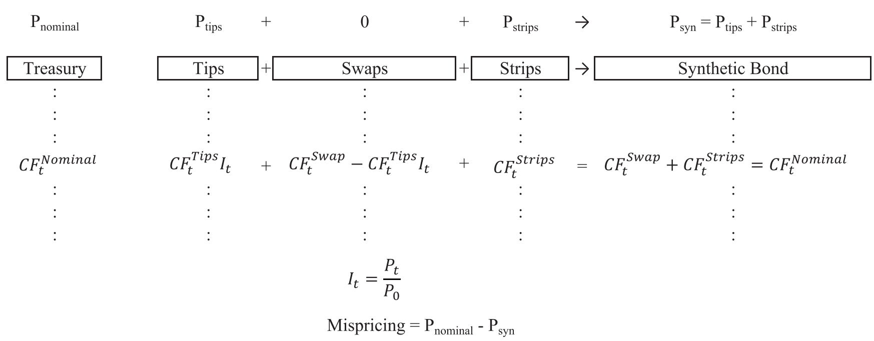

Fig A1. Arbitrage Strategy.

NOTE: This figure demonstrates how one can replicate the cash flow from a nominal bond using inflation linked bonds, inflation swaps, and strips. When the trade is initialized at time zero, the cash flow on an inflation linked bond that will arrive at time t (CF_t^{Tips}) can be converted into a nominal cash flow CF_t^{Swap} by entering into an inflation swap at time 0. By purchasing a zero coupon bond (strip) at time zero the arbitrageur can insure receipt of a cash flow CF_t^{Strips} that equals the difference between the cash flow on the nominal bond ($CF_t^{Nominal}$) and the nominal cash flow from the inflation swap. We then have two securities with the same payoffs. The first is a nominal sovereign bond. The second is a synthetic bond made up of the inflation linked bond, zero coupon strips, and inflation swaps. The arbitrageur then purchases the cheaper bond and sells the more expensive bond. This strategy ensures a positive profit today with zero net cost for all future periods.

TABLE A1

BLOOMBERG TICKERS FOR VARIABLES USED

Ticker	Name
HFRXGL Index	Global Hedge Fund Index
CDX IG CDSI GENERIC 5Y Corp	CDX North America Investment Grade Index
ITRAX JAPAN CDSI GENERIC 5Y Corp	iTraxx Japan Investment Grade Index
ITRX EUR CDSI GENERIC 5Y Corp	iTraxx Europe Investment Grade Index
CONCCONF Index	Consumer Confidence Index US
JCOMACF Index	Japan Consumer Confidence Nationwide NSA
UKCCI Index	GFK UK Consumer Confidence Indicator
GRCCI Index	ICon Germany Consumer Confidence Indicator
FRCCO Index	France Consumer Confidence Indicator SWDA
ITPSSA Index	Italy Consumer Confidence Indicator SA
VIX Index	Chicago Board Options Exchange Volatility Index
VXJ Index	Nikkei Stock Average Volatility Index
VFTSE Index	FTSE 100 Volatility Index
V2X Index	Euro Stoxx 50 Volatility Index VSTOXX
HFRXGL Index	HFRX Global Hedge Fund Index
HFRXMMS Index	HFRX Macro: Multi-Strategy Index

(*Continued*)

TABLE A1

(Continued)

Ticker	Name
HFRXRVA Index	HFRX Relative Value Arbitrage Index
HFRXCA Index	HFRX RV: FI-Convertible Arbitrage Index
HFRXCA Index	HFRX RV: FI-Convertible Arbitrage Index
HFRXVOL Index	HFRX RV: Volatility Index
HFRXYA Index	HFRX RV: Yield Alternative Index
HFRXFSV Index	HFRX RV: FI-Sovereign Index
SP Govt	United States Treasury Strip Principal
JGBS Govt	Japan Government Coupon Strips
UKTS Govt	United Kingdom Gilt Coupon Strip
DBRS Govt	Deutsche Bundesrepublik Coupon Strips
FRTRR Govt	France Government Bond OAT Principal Strip
BTPSS Govt	Italy Buoni Poliennali del Tesoro Coupon Strip
USSWIT1 Curncy	USD INFLATION SWAP ZERO CPN 1Y
JYSWIT1 CMPL Curncy	JY INFLATION SWAP ZERO COUPON 1Y
BPSWIT1 Curncy	UK INFLATION SWAP ZERO COUPON 1Y
EUSWIT1 Curncy	EURO INFLATION SWAP ZERO CPN 1Y
US CDS USD SR 5Y Corp	USA Govt 5-year CDS Contract
JGB CDS USD SR 5Y Corp	Japan Govt 5-year CDS Contract
UK CDS USD SR 5Y Corp	UK Govt 5-year CDS Contract
GERMAN CDS USD SR 5Y Corp	German Govt 5-year CDS Contract
FRANCE CDS USD SR 5Y Corp	France Govt 5-year CDS Contract
ITALY CDS USD SR 5Y Corp	Italy Govt 5-year CDS Contract

TABLE A2

Summary Statistics for Mispricing (Bond Level)

Country	Bond	Mean	Std Dev	Min	Max	Maturity Mismatch
United States	9128273T7	0.474	0.608	–0.758	1.569	16
	9128274Y5	1.150	1.997	–2.305	5.835	16
	9128276R8	–0.497	3.366	–6.755	12.923	0
	9128277J5	–8.578	2.435	–16.267	–2.417	0
	912828HW3	0.614	0.721	–1.264	2.244	0
	912828BW9	0.132	1.766	–3.470	3.220	16
	912828KM1	0.650	1.012	–2.252	1.996	0
	912828DH0	0.169	1.025	–2.405	2.347	16
	912828MY3	0.686	0.466	–0.747	1.994	15
	912828ET3	0.918	1.512	–2.518	3.768	16
	912828FL9	–0.178	1.320	–3.023	2.451	16
	912828GD6	0.428	1.757	–3.454	3.837	16
	912828GX2	–1.063	1.996	–4.423	2.655	16
	912828HN3	1.778	2.192	–2.258	15.230	0
	912828JE1	2.794	2.424	–0.163	16.210	31
	912828JX9	0.295	2.298	–4.192	5.142	31
	912828LA6	0.366	1.325	–2.408	3.418	31
	912828MF4	1.916	1.079	–0.573	4.460	31
	912828NM8	5.652	2.960	–0.687	13.116	31
	912828PP9	4.275	1.667	–0.153	7.956	31
	912828QV5	3.787	1.937	–0.811	9.166	31

(*Continued*)

TABLE A2
(CONTINUED)

Country	Bond	Mean	Std Dev	Min	Max	Maturity Mismatch
	912810FR4	2.764	14.767	–23.562	43.375	31
	912810FS2	3.271	9.306	–15.733	35.995	31
	912810PS1	7.256	5.052	–3.908	29.410	31
	912810PZ5	3.887	2.928	–1.216	15.289	31
United Kingdom	ED361990	1.872	3.336	–6.497	13.976	89
	ED970564	1.858	1.052	–0.165	5.011	15
	EF315225	1.262	0.740	–0.364	2.716	168
	EF2659706	10.299	6.158	–1.113	28.036	169
	EF372237	1.755	3.798	–8.866	17.008	15
Japan	EH600918	–0.380	3.569	–7.131	12.071	66
	EI684934	–6.576	4.694	–16.622	3.243	80
	EG196397	–1.278	4.690	–11.945	14.961	80
Germany	EF3134212	4.621	2.158	0.562	9.281	172
	EI639514	2.873	0.995	0.992	6.338	41
	EH8565820	1.902	1.492	–1.535	6.147	10
	EJ0993182	1.646	1.426	–1.436	7.436	10
France	EI540344	2.746	1.369	–0.139	6.392	91
	EF081090	2.934	2.262	–0.617	12.295	92
	EI112670	0.895	1.296	1.296	–2.022	92
	EH212767	0.818	2.262	–4.266	12.465	91
	EC182706	–1.338	3.420	–9.982	15.764	91
Italy	ED327992	1.533	1.508	–2.244	7.279	45
	EI548734	3.810	1.963	–0.869	9.560	45
	EF504151	5.066	2.674	–2.215	18.467	45
	EH378395	3.018	2.803	–1.963	15.426	14
	EI230886	8.879	2.921	3.183	17.148	45

NOTE: This table reports the summary statistics for the dollar-index and nominal bond mispricing for the 47 pairs of six G7 countries. The mispricing is measured in dollars per $100 notional. The last column reports the numbers of days of the maturity mismatch of our pairs, as discussed in Section 2. The sample period spans from February 2, 2007, to November 30, 2012.

APPENDIX B: STRUCTURAL VAR

Let $q_t = \{m_t, i_t, ip_t, oil_t\}$ be the vector of variables: the *mispricing*, *overnight interest rate*, *industrial production*, and *oil price*. We can write the VAR in matrix form:

$$q_t = Bq_{t-1} + u_t \text{ with } Var(u_t) = \Sigma,$$

where u_t are the forecast errors of the VAR. We assume there are four structural disturbances in the economy $\varepsilon_t = ({\varepsilon_t}^1 {\varepsilon_t}^2 {\varepsilon_t}^3 {\varepsilon_t}^4)'$ with $Var(\varepsilon_t) = I$, which relate to the forecast errors by $u_t = S\varepsilon_t$. Here ε_t^i is the shock to the short-term interest rate. We assume that S is a lower triangular matrix. This implies that the shocks to the short-term interest rate do not effect *industrial production* and *oil prices* within the month and that the central bank does not respond to the *mispricing* in setting interest rates. We then have

$$\Sigma = SS'$$

and therefore S can be recovered using the Cholesky decomposition on the estimate of the forecast error variance–covariance matrix. Finally we can calculate impulse responses using the dynamic system setting $q_0 = \overrightarrow{0}$

$$q_t = Bq_{t-1} + S\varepsilon_t.$$

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SUPPORTING INFORMATION

Additional supporting information may be found online in the Supporting Information section at the end of the article.

Data S1