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Jens H. E. Christensen
Federal Reserve Bank of San Francisco

Nikola Mirkov
ITAM

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The Safety Premium of Safe Assets

Jens H. E. Christensen[†]

&

Nikola Mirkov[‡]

Abstract

Safe assets usually trade at a premium thanks to high credit quality and deep liquidity. To understand the role of credit quality for such premia, we examine Swiss government bonds, which are extremely safe but not particularly liquid. We therefore refer to their premia as *safety premia* and quantify them using an arbitrage-free term structure model that accounts for time-varying premia in individual bond prices. We find that Swiss safety premia are large with long-lasting trends. They shifted upwards persistently following the euro launch and have been depressed recently by asset purchases of the European Central Bank.

JEL Classification: C32, E43, E52, F34, G12

Keywords: affine arbitrage-free term structure model, bond-specific risk premia, euro launch, negative interest rates

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[†]Corresponding author: Federal Reserve Bank of San Francisco, 101 Market Street MS 1130, San Francisco, CA 94105, USA; phone: 1-415-974-3115; e-mail: jens.christensen@sf.frb.org.

[‡]Visiting researcher, ITAM; e-mail: nikola.mirkov@gmail.com.

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1 Introduction

A growing body of research has emphasized the importance of safe assets for a number of global macroeconomic trends. Due to global investment and savings imbalances between developed and emerging market economies,[1] as argued by Caballero et al. (2017), a shortage of safe assets partially explains the low level of interest rates in many countries.[2] Furthermore, the shortage of safe assets has implications for financial stability if it forces investors to search for higher yields, as discussed in Gorton (2017). Finally, since nominal yields are constrained by a lower bound near zero, a lack of safe assets may also be one of the factors that has constrained monetary policy around the world following the global financial crisis.[3] Hence, the properties of safe assets and their pricing merit further examination.

Previous research in this area has focused mainly on U.S. Treasury securities and documented the existence of a premium in their prices arising from a combination of high credit quality and deep liquidity.[4] Such premia are typically referred to as convenience premia and defined as the difference between the observed yield and the fundamental yield that the assets would pay without any value attached to their special attributes. The convenience premium can be broken down further into a liquidity premium and a safety premium. The liquidity premium represents the value investors are willing to pay in exchange for the ability to easily buy and sell the asset. The safety premium is the value investors attach to the safety of the asset, i.e., the property that the asset is likely to hold up its value even under severely adverse financial market conditions.

In this paper, we examine the convenience premium embedded in the prices of an asset widely viewed as one of the safest government bonds in the world, namely Swiss government bonds, also known as Confederation bonds or Eidgenossen in German. Given that these bonds are far less liquid than their U.S. counterparts, the convenience premia of Swiss Confederation bonds can be almost entirely attributed to their high credit quality.[5] We therefore refer to them as *safety premia* and measure them as the lower yield earned on these assets relative to the estimated yields of otherwise identical assets with the same coupon and principal payments.

Our focus on Swiss Confederation bonds has several additional motivations. First, few papers have estimated either liquidity or safety premia for standard government bond markets outside of the U.S. Treasuries market. Thus, little is known about the magnitudes of such premia in regular sovereign bond markets.[6]

[1]Bernanke (2005) is frequently referenced regarding this topic.

[2]Glick (2019) provides an excellent overview of the recent literature on this topic with a global perspective.

[3]See Reifschneider and Williams (2000) for an early discussion of constraints on monetary policy in low-interest-rate environments.

[4]Early examples of papers that estimate the liquidity convenience yields of U.S. Treasuries include Amihud and Mendelson (1991) and Longstaff (2004), among many others.

[5]The high safety of these bonds is in part ensured by a debt ceiling limit for the Swiss government in much the same way as the one that applies to the U.S. federal government.

[6]Christensen et al. (2021) apply an approach similar to ours to estimate illiquidity premia in the sovereign

Second, given that the Swiss National Bank (SNB) was one of the first central banks to introduce negative policy rates and has gone further in that direction than any of its peers, Switzerland offers a unique case for studying the behavior of safety premia in various regimes: (1) the normal period before the financial crisis; (2) the 2008-2014 period in which the Swiss monetary policy rate was cut decisively but remained in positive territory; and (3) the 2015-2019 period in which Swiss interest rates moved deeply into negative territory after the SNB discontinued its Swiss franc to euro minimum exchange rate. Specifically, this latter episode allows us to analyze whether safety premia are affected when yields turn negative.

Last, our data also allow us to study whether asset purchases by the European Central Bank (ECB) have affected Swiss safety premia. Given that such purchases globally reduce the available stock of safe assets, they could increase the safety premia in many countries, including Switzerland. Alternatively, in a narrow European context, such purchases reduce the stock of euro-area safe assets. This reduction makes Swiss safe assets less exclusive relative to those of its European neighbors, potentially lowering the safety premium that Swiss safe assets can require.[7] Ultimately, which of these effects dominates in the pricing of Swiss safe assets is an empirical question.

To quantify the safety premia in the Swiss Confederation bond market, we use the approach of Andreasen et al. (2021, henceforth ACR), who augment a standard term structure model with a risk factor to accurately measure bond-specific risk premia. In addition to assuming a common component in the bond-specific risk premia captured by the added risk factor, the model stipulates that such premia are a function of time since issuance and remaining time to maturity. This is intended to mimic the idea that the excess premium investors are willing to pay for a safe bond depends on the amount of it available for trading. As time passes, this amount tends to decline as an increasing fraction of its outstanding notional amount gets locked up in buy-and-hold investors' portfolios.[8] This raises its sensitivity to variation in the market-wide systematic component of the bond-specific safety premium captured by the added risk factor. By observing the entire cross section of government bond prices over time, this factor can be separately identified from the conventional fundamental risk factors in the model.

Since Swiss Confederation bonds are widely considered not to be particularly liquid and may carry an outright liquidity discount for that reason, one could be concerned that our estimated bond-specific safety premia might be polluted by such liquidity effects. To address that issue, we implement an expanded version of our baseline model that includes a separate liquidity risk factor in addition to the bond-specific risk factor described above and find the results to be robust to such extensions.

bond market of Mexico, a major emerging economy.

[7]Krishnamurthy and Vissing-Jorgensen (2012) find that a lower supply of U.S. Treasury debt widens the U.S. Treasury yield spread relative to AAA-rated corporate bonds. We therefore hypothesize that a similar effect could be present in the relative pricing between safe assets in the euro-area and Swiss safe assets.

[8]Swiss Confederation bonds fit this characterization as they are not viewed as particularly liquid.

Our results can be summarized as follows. First, we find that the augmented model improves model fit and delivers robust estimates of the conventional risk factors that drive the variation in the frictionless part of the Swiss Confederation bond yield curve. Furthermore, the fit is not affected by either the 2008-2014 period near the zero lower bound (ZLB) or the negative rate environment prevailing since December 2014. Actually, the fit is better in these periods than during the first 15 years of our sample, with interest rates far above zero and more variable.

Second, our estimation results show that the safety premium has been positive on average, with a mean of 68 basis points and a standard deviation of 20 basis points. This likely reflects the flight-to-safety advantages Swiss government-backed securities offer investors.[9] Their sizable mean and variability also underscore that safety premia are important components in the pricing of Swiss Confederation bonds, in particular with Swiss interest rates near historical lows recently.

Most interestingly, the introduction of the euro in January 1999 appears to have given rise to an upward shift in the Swiss safety premium. A regression analysis with a variety of control variables suggests that the premium experienced a long-lasting spike of between 35 and 40 basis points around this event. We conjecture that this shift might be caused by the perceived implicit risk-sharing across eurozone countries that could have reduced the safety of government bonds in core eurozone countries such as Germany or France relative to that of Swiss government bonds.

Furthermore and relatedly, our regression results indicate that the ECB's purchases of government bonds since January 2015 have exerted a persistent downward pressure on the Swiss safety premium with a cumulative effect of about 20 basis points.[10] In contrast, we find that the introduction of negative interest rates by the SNB cannot explain the gradual decline of the average safety premium since late 2014, once we control for all confounding factors. Hence, our results suggest that the scarcity of euro-area safe assets created by the ECB's asset purchases might have reduced the relative exclusiveness of Swiss safe assets and thereby depressed their safety premium.

Finally, as an additional contribution, we exploit our yield curve model to study the term structure of safety premia. The results show that there was an upward slope in the premia in the 1990s, followed by a relatively flat term structure in the period from the launch of the euro until the global financial crisis. Since 2009, however, the term structure has turned negative, driven by a much sharper decline in the long-term safety premia than in the medium-term safety premia. We speculate that this may be tied to the very high duration risk of long-term bonds in low-interest-rate environments and could be an indication that

[9] Consistent with our results, Jäggi et al. (2016) describe flight-to-safety effects in the Swiss franc.

[10] Note that this effect on convenience yields of safe assets is different from the international spillover effects of central bank asset purchases, such as the signaling or portfolio balance effects traditionally discussed in the literature, which affect the general expectations and risk premium components of bond yields; see Bauer and Neely (2014) for an example.

even the most desirable safe assets can lose some of their attractiveness when interest rates become sufficiently low.

Before drawing any conclusions based on the apparent connection between the lower trend in the safety premium and the coincidental decline in longer-term Swiss interest rates, note that these changes happened in the context of relatively low financial market volatility, at least compared to that in both the global financial crisis of 2008-2009 and the subsequent European sovereign debt crisis in 2011-2012. Thus, to what extent flight-to-safety effects, which tend to spike when economic uncertainty is elevated, are truly smaller in low-interest-rate environments is difficult to determine based on our data.

As for the underlying cause for the existence of the safety premia we identify in Swiss Confederation bond prices, these premia may be related to the "safe haven" status of the Swiss franc, which is typically taken to mean that safe Swiss franc-denominated assets offer a hedging value against global risk, both on average and specifically in crisis episodes; see Grisse and Nitschka (2015) for evidence and a discussion.

In a final exercise, we repeat our analysis using Danish government bonds. Similar to Swiss Confederation bonds, these bonds are widely viewed as being among the safest government bonds in the world but lack the liquidity superiority of U.S. Treasuries. Therefore, we consider any bond-specific premia in their prices to represent safety premia. The regression analysis using Danish government bonds produces results very similar to those derived from the Swiss Confederation bonds. First and foremost, the launch of the euro appears to have boosted Danish safety premia by 20-30 basis points. Second, Danish safety premia were not affected by the introduction of negative rates by the Danish National Bank (DNB) in July 2012. Finally, the ECB's purchases of safe government bonds from the euro area have also put significant downward pressure on the Danish safety premia.

Combined our findings point to a potentially important international spillover channel of central bank asset purchases that operates through its impact on the relative scarcity of safe assets.

The remainder of the paper is organized as follows. Section 2 provides a brief summary of various definitions of safe assets and puts our identification approach into context. Section 3 describes the Swiss Confederation bond data, while Section 4 details the no-arbitrage term structure model we use and presents the empirical results. Section 5 analyzes the estimated Confederation bond safety premium and its determinants, while Section 6 is dedicated to a similar analysis of the estimated Danish safety premia. Section 7 concludes. An online appendix contains details of the data and a description of our control variables.

2 Related Literature on Safe Assets

To better understand the market pricing of safe assets, it is useful to first discuss what makes an asset "safe." In the existing literature, there are many overlapping definitions. Gorton

(2017) describes a safe asset as one that is (almost always) valued at face value without expensive and prolonged analysis, i.e., there is little to no value of private information and hence no risk of selection bias or other strategic motives for trading. Gourinchas and Jeanne (2012) define it as a liquid debt claim with negligible default risk that provides a secure store of value. Caballero et al. (2017) emphasize that it is a debt instrument that is expected to preserve its value during adverse systemic events. Hence, common threads across these definitions are that a safe asset should have the properties of providing security, such that it pays close to par with near certainty in the future, and providing liquidity, such that it is similar to money in its availability and acceptability.

As a consequence of these characteristics, safe assets serve as a useful benchmark for several important tasks. They play a transactional role by serving as collateral in financial transactions and regulatory capital in meeting liquidity requirements. They serve as an accessible store of value by providing a reliable return. In addition, they serve an accounting role as a benchmark for the pricing of other assets (see IMF 2012). In light of these useful attributes, we should expect the value of these services to be reflected in the prices of safe assets.

In terms of this question, previous research has focused mainly on U.S. Treasury securities and documented the existence of a premium in their prices arising from a combination of high credit quality and deep liquidity. While it may be possible to estimate the convenience premium of a safe asset, it is generally challenging to separate it into the underlying liquidity and safety premia. To achieve a better understanding of each of these components, Krishnamurthy and Vissing-Jorgensen (2012) scrutinize the yield spreads of assets with differences in liquidity but similar safety and those of assets with different levels of safety but similar liquidity. They find that U.S. Treasury yields averaged 73 basis points lower than they otherwise would have during the 1926-2008 period due to their safety and liquidity. Using the yield spreads of AAA- and BBB-rated corporate bonds, their analysis further suggests that up to 46 basis points of this convenience yield reflect the liquidity advantage of U.S. Treasuries, which leaves at least 27 basis points to be explained by their safety premium.

Nagel (2016) focuses on matching the three-month Treasury bill and repo rates. Given that the latter is equivalent to a collateralized loan, these are both claims that are essentially free of credit risk. Hence, the Treasury bills' convenience yield averaging 24 basis points in the 1991-2011 period should exclusively represent liquidity premia arising from their extreme liquidity. Combined with the results from Krishnamurthy and Vissing-Jorgensen (2012), this would leave a safety premium of about 50 basis points in U.S. Treasury yields.

In this paper, we follow an alternative approach and analyze assets that are extremely safe but relatively illiquid. By implication, any convenience yield in their pricing must mainly reflect a safety premium and is unlikely to represent any liquidity premium. This should allow us to shed light on the role of credit quality and safety for the convenience premia of safe assets.

2.1 The Identification of Safety Premia

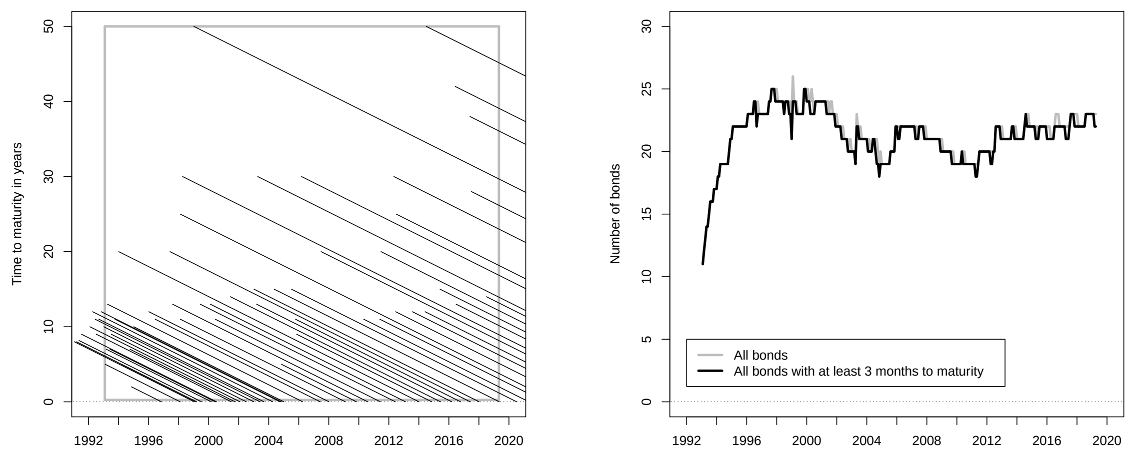
As described earlier, our model discounts coupon and principal payments from individual Confederation bonds by using bond-specific short rates, which differ in their loadings on the common bond-specific risk factor. This is a very direct way of identifying bond-specific risk premia and requires only a panel of bond market prices, which is readily available. A key advantage of our approach is that we do not need any data on bond-specific trading dynamics such as bid-ask spreads, trading volumes, quote sizes, or trade sizes to be able to identify the unique premia embedded in individual bond prices. Furthermore, as emphasized by ACR, such measures reflect *current* market conditions, whereas our concept of bond-specific safety premia is centered around investors' expectations about *future* market conditions and how those expectations should affect *current* bond prices.

A theoretical implication of our model is that bond-specific safety premia are identified from the implied price differential of otherwise identical principal and coupon payments. The proposed identification scheme is thus related to the analysis of Fontaine and Garcia (2012). However, unlike that study, our approach does not require the existence of nearly matching pairs of securities for identification and hence can be applied to any bond market. Furthermore, our model and its application differ along other dimensions from the analysis in Fontaine and Garcia (2012). In particular, we impose the absence of arbitrage and hence allow the bond-specific safety premium adjustment to factor into the expectations formed by investors.

Furthermore, provided Swiss Confederation bonds are considered to be illiquid and carry an outright liquidity discount, our estimated safety premia could be viewed as lower bound estimates of the underlying true safety premium. Thus, we consider our analysis conservative from this perspective.

3 Swiss Confederation Bond Data

Our Swiss Confederation bond price data are collected daily by staff at the SNB and are available back to the 1980s. However, an inspection of the data reveals that the early part of the sample is characterized by many stale or erratic prices. For these reasons we choose to start the data sample in January 1993, a date at which the data appear to be systematically reliable across all available bonds. For the purposes of factor identification, the model needs at least four bonds to be traded on each day in the sample. As a consequence, we track a few select bonds that were issued before January 1993 and appear to have reliable prices. These considerations lead us to focus on the universe of Swiss Confederation bonds listed in online



(a) Maturity distribution of Confederation bonds

(b) Number of Confederation bonds

Figure 1: **The Swiss Confederation Bond Market**
Panel (a) shows the maturity distribution of the Swiss Confederation bonds considered. The solid gray rectangle indicates the subsample used throughout the paper. Panel (b) reports the number of Swiss Confederation bonds at each date.

Appendix A.[11] Importantly, the sample contains every Swiss Confederation bond issued since 1993. Thus, our analysis is complete and comprehensive for the period since then. While the bond data are available at a daily frequency, the analysis was performed on monthly data to facilitate empirical implementation.

Figure 1(a) shows the maturity distribution of the universe of Swiss Confederation bonds across time since 1993. The vertical solid gray lines indicate the start and end dates for our sample, while the horizontal solid gray lines indicate the top and bottom of the maturity range considered. The top of the range equals 50 years and is determined by the longest bond maturity issued by the Swiss Confederation, while the bottom of the range is fixed at three months to avoid erratic price patterns for bonds as they approach maturity, see Gürkaynak et al. (2007).

Figure 1(b) shows the number of bonds in our sample at each point in time. We note that the total number of outstanding bonds has been fairly stable since the mid-1990s. At the end of our sample period, a total of 22 bonds remained outstanding.

Figure 2 shows the Swiss Confederation bond prices converted into yield to maturity. Note the following regarding these yield series. First, yield levels have generally trended lower during this 26-year period, ranging from close to 5 percent in the early 1990s to zero by the end of our sample. Second, business cycle variation in the shape of the yield curve is pronounced around the lower trend. The yield curve tends to flatten ahead of recessions and steepen during the initial phase of economic recoveries. It is these characteristics that are the

[11] Our sample represents an update of the data used by Christensen at al. (2021).

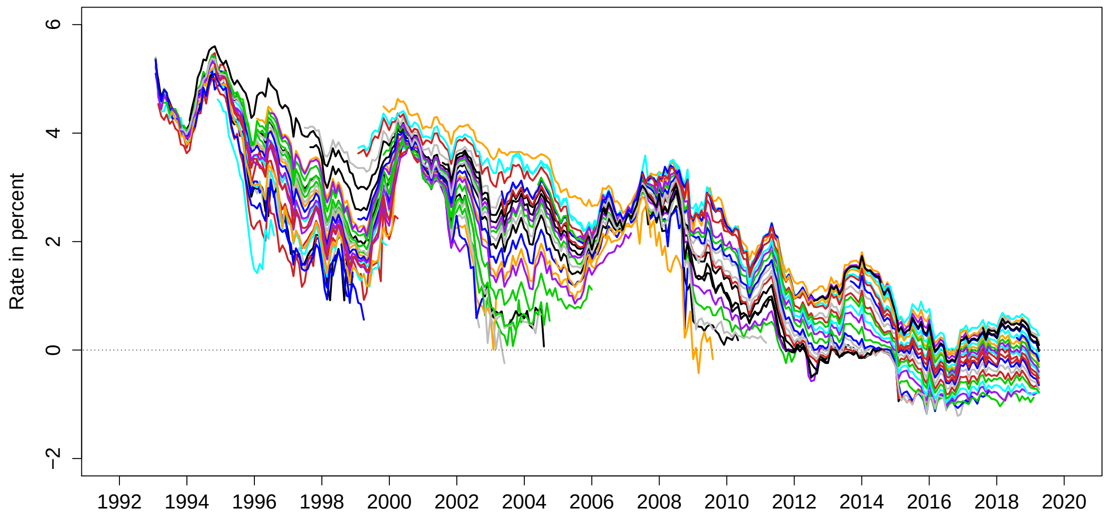


Figure 2: **Yield to Maturity of Swiss Confederation Bonds**
Illustration of the yield to maturity of the Swiss Confederation bonds considered in this paper, which are subject to two sample choices: (1) sample limited to the period from January 29, 1993, to March 29, 2019; (2) censoring of a bond's price when it has less than three months to maturity.

practical motivation behind our choice of using a three-factor model for the frictionless part of the Swiss yield curve, adopting an approach similar to what is standard for U.S. and U.K. data; see Christensen and Rudebusch (2012).

To support that choice more formally, we note that researchers have typically found that three factors are sufficient to model the time variation in the cross section of U.S. Treasury yields (e.g., Litterman and Scheinkman 1991). To perform a similar analysis based on our sample of Swiss Confederation bond prices, we construct synthetic zero-coupon bond yields. For the period from January 1993 to December 1997 with a limited range of available bond maturities, we follow De Pooter et al. (2014) and use the Nelson and Siegel (1987) yield curve function for this construction, while we switch to the more flexible Svensson (1995) yield curve for the period since January 1998 due to the greater dispersion in the available set of bond maturities.[12] To have a yield panel representative of the underlying bonds in our sample, we include yields for seven constant maturities: 2, 3, 5, 7, 10, 12, and 20 years. The data series are daily series, covering the period from January 4, 1993, to March 29, 2019.

The result of a principal component analysis of the yield panel is reported in Table 1. The top panel reports the eigenvectors that correspond to the first three principal components. The first principal component accounts for 93.2 percent of the variation in the bond yields, and its loading across maturities is uniformly positive. Thus, similar to a level factor, a

[12] Technically, for both periods, we proceed as described in Andreasen et al. (2019).

Maturity in months	First P.C.	Second P.C.	Third P.C.
24	0.42	0.84	-0.24
36	0.39	0.14	0.45
60	0.38	-0.15	0.46
84	0.37	-0.21	0.24
120	0.36	-0.24	-0.06
144	0.36	-0.25	-0.22
240	0.36	-0.28	-0.65
% explained	93.21	5.83	0.79

Table 1: **Factor Loadings of Swiss Confederation Bond Yields**
The top rows show the eigenvectors corresponding to the first three principal components (PC). Put differently, they show how bond yields at various maturities load on the first three principal components. In the final row the proportion of all bond yield variability explained by each principal component is shown. The data are daily Swiss Confederation zero-coupon bond yields from January 4, 1993, to March 29, 2019, a total of 6,578 observations for each yield series.

shock to this component changes all yields in the same direction irrespective of maturity. The second principal component accounts for 5.8 percent of the variation in these data and has sizable positive loadings for the shorter maturities and sizable negative loadings for the long maturities. Thus, similar to a slope factor, a shock to this component steepens or flattens the yield curve. Finally, the third component, which accounts for 0.8 percent of the variation, has a hump shaped factor loading as a function of maturity, which is naturally interpreted as a curvature factor. These three factors combined account for 99.83 percent of the total variation. This motivates our choice to focus on the Nelson and Siegel (1987) model with its level, slope, and curvature factors for modeling this sample of Swiss bond yields. However, for theoretical consistency, we use the arbitrage-free version of this class of models derived in Christensen et al. (2011). Furthermore, to explain the remaining variation in the bond yield data not accounted for by the level, slope, and curvature factors, we augment the model with a risk factor to allow for bond-specific risk premia using the approach described in ACR and detailed in Section 4. Importantly, we stress that the estimated state variables in our model are *not* identical to the principal component factors discussed here, but estimated through Kalman filtering.[13]

Finally, we note that Swiss yields had come close to the ZLB already in 2003-2004. Thus, Swiss fixed-income markets have a long history of being at or near the ZLB, only matched by their Japanese equivalents. However, unlike Japanese yields that mostly respected the ZLB until 2015 (see Christensen and Spiegel 2021), Swiss yields have been less constrained by the ZLB. Considering the data available in 2019 when several yields were close to negative 1 percent, it is not clear what the lower bound for Swiss yields is. As a consequence, we

[13]A number of recent papers use principal components as state variables. Joslin et al. (2011) is an early example.

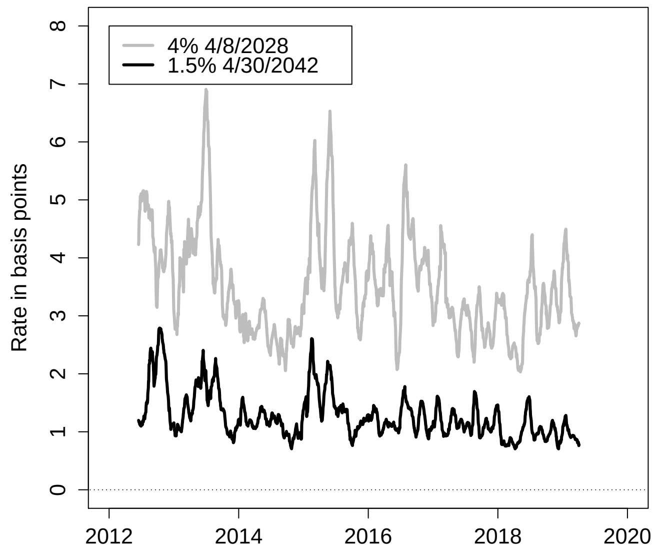


Figure 3: **Bid-Ask Spreads of Thirty-Year Swiss Confederation Bonds**
Illustration of the four-week moving average of bid-ask spreads of two thirty-year Swiss Confederation bonds. The series are daily covering the period from June 12, 2012, to March 29, 2019.

choose to focus on models with Gaussian dynamics, which can easily handle negative interest rates.[14]

3.1 The Bond-Specific Risk of Confederation Bonds

In this paper, we apply the ACR model approach to the universe of Swiss Confederation bonds. To support the use of the ACR approach as a way of identifying bond-specific risk premia in Confederation bond prices, we point to the structure of bonds' bid-ask spreads. Figure 3 shows two series of bid-ask spreads for thirty-year Confederation bonds, one represents the bid-ask spread of the first thirty-year bond in our sample (4 percent Eidgenosse maturing 4/8/2028), the other tracks the bid-ask spread of the most recently issued thirty-year bond in our sample (1.5 percent Eidgenosse maturing 4/30/2042) issued in April 2012. Both series are smoothed four-week averages and are measured in basis points. Similar to the findings of ACR for U.S. Treasury inflation-protected securities (TIPS), the pattern in the bid-ask spreads of Swiss Confederation bonds is such that the more seasoned securities are less liquid than the recently issued securities. Rational, forward-looking investors are aware of these dynamics and the fact that future market liquidity of a given security is likely to be below its current

[14]This choice is also supported by the analysis of Grisse and Schumacher (2018), who find that, with the exception of a brief period, long-term Swiss Confederation bond yields have responded symmetrically to changes in the short rate since 2000, a pattern well captured by Gaussian models.

market liquidity. These patterns give rise to security-specific premia in the bond prices.[15] The ACR approach detailed below is designed to capture such premia in individual bond prices. Also, these series underscore the relatively low liquidity of seasoned Confederation bonds, which dominate our sample.

4 Model Estimation and Results

In this section, we detail the term structure model that serves as the benchmark in our analysis and describe the restrictions imposed to achieve econometric identification of the model. We then compare its estimates to those from the model without any adjustment for bond-specific risk premia. Finally, we provide a more detailed analysis of the fit of the model relative to a set of relevant benchmark models.

4.1 The AFNS-R Model

The fundamental frictionless Swiss yields that would prevail in a world without any frictions to trading or any excess demand attributable to safety concerns regarding other assets are modeled by using a standard Gaussian model, namely the arbitrage-free Nelson-Siegel (AFNS) model introduced in Christensen et al. (2011). We augment this model with a risk factor structured as in the ACR approach and refer to it as the AFNS-R model.

To begin the model description, let $X_t = (L_t, S_t, C_t, X_t^R)$ denote the state vector of the four-factor AFNS-R model. Here, L_t denotes a level factor, while S_t and C_t represent slope and curvature factors. Finally, X_t^R is the added marketwide bond-specific risk factor.

The instantaneous risk-free rate is defined as follows:

$$r_t = L_t + S_t. \tag{1}$$

The risk-neutral dynamics of the state variables used for pricing are given by the following:

$$\begin{pmatrix} dL_t \\ dS_t \\ dC_t \\ dX_t^R \end{pmatrix} = \begin{pmatrix} 0 & 0 & 0 & 0 \\ 0 & \lambda & -\lambda & 0 \\ 0 & 0 & \lambda & 0 \\ 0 & 0 & 0 & \kappa_R^{\mathbb{Q}} \end{pmatrix} \left[\begin{pmatrix} 0 \\ 0 \\ 0 \\ \theta_R^{\mathbb{Q}} \end{pmatrix} - \begin{pmatrix} L_t \\ S_t \\ C_t \\ X_t^R \end{pmatrix} \right] dt + \Sigma \begin{pmatrix} dW_t^{L,\mathbb{Q}} \\ dW_t^{S,\mathbb{Q}} \\ dW_t^{C,\mathbb{Q}} \\ dW_t^{R,\mathbb{Q}} \end{pmatrix},$$

where Σ is a lower-triangular matrix.

Based on the $\mathbb{Q}$-dynamics above, the standard frictionless zero-coupon bond yields preserve a Nelson and Siegel (1987) factor loading structure:

$$y_t(\tau) = L_t + \left(\frac{1 - e^{-\lambda\tau}}{\lambda\tau} \right) S_t + \left(\frac{1 - e^{-\lambda\tau}}{\lambda\tau} - e^{-\lambda\tau} \right) C_t - \frac{A(\tau)}{\tau}, \tag{2}$$

[15] Fontaine and Garcia (2012) document pricing differences of this nature in the U.S. Treasury market.

where $\frac{A(\tau)}{\tau}$ is a convexity term that adjusts the functional form in Nelson and Siegel (1987) to ensure absence of arbitrage (see Christensen et al. 2011).

Importantly, due to bond-specific premia in the Confederation bond market, individual bond prices are sensitive to the variation in the common component of the bond-specific premia captured by the added risk factor X_t^R. As a consequence, the pricing of each Confederation bond is not performed with the standard discount function, but rather with a discount function that accounts for its bond-specific risk:

$$\overline{r}_t^i = r_t + \beta^i (1 - e^{-\lambda^{R,i}(t - t_0^i)}) X_t^R, \tag{3}$$

where t_0^i denotes the date of issuance of the specific security and β^i is its sensitivity to the variation in the associated market-wide risk factor. Furthermore, the decay parameter $\lambda^{R,i}$ is assumed to vary across securities as well.

As shown in Christensen and Rudebusch (2019), the net present value of one Swiss franc paid by Confederation bond i at time $t + \tau$ has the following exponential-affine form:

$$\begin{aligned} P_t^i(t_0^i, \tau) &= E^{\mathbb{Q}}\left[e^{-\int_t^{t+\tau} \overline{r}^i(s, t_0^i) ds}\right] \\ &= \exp\left(B_1^i(\tau) L_t + B_2^i(\tau) S_t + B_3^i(\tau) C_t + B_4^i(t_0^i, t, \tau) X_t^R + A^i(t_0^i, t, \tau)\right). \end{aligned}$$

This implies that the model belongs to the class of Gaussian affine term structure models. Note also that, by fixing $\beta^i = 0$ for all i, we recover the AFNS model.

Consider the whole value of the Swiss Confederation bond issued at time t_0^i with maturity at $t + \tau$ that pays a coupon C annually. Its price is given by the following:[16]

$$P_t^i(t_0^i, \tau) = C(t_1 - t) E^{\mathbb{Q}}\left[e^{-\int_t^{t_1} \overline{r}^i(s, t_0^i) ds}\right] + \sum_{j=2}^{N} C E^{\mathbb{Q}}\left[e^{-\int_t^{t_j} \overline{r}^i(s, t_0^i) ds}\right] + E^{\mathbb{Q}}\left[e^{-\int_t^{t+\tau} \overline{r}^i(s, t_0^i) ds}\right]. \tag{4}$$

At present, the description of the AFNS-R model has relied solely on the dynamics of the state variables under the $\mathbb{Q}$-measure used for pricing. However, to complete the description of the model and to implement it empirically, we will need to specify the risk premia that connect these factor dynamics under the $\mathbb{Q}$-measure to the dynamics under the real-world (or physical) $\mathbb{P}$-measure. Note that there are no restrictions on the dynamic drift components under the empirical $\mathbb{P}$-measure beyond the requirement of constant volatility. To facilitate empirical implementation, we use the essentially affine risk premium specification introduced in Duffee (2002). In the Gaussian framework, this specification implies that the risk premia Γ_t depend on the state variables; that is,

$$\Gamma_t = \gamma^0 + \gamma^1 X_t,$$

[16] This is the clean price that does not account for any accrued interest and maps to our observed bond prices.

where $\gamma^0 \in \mathbf{R}^4$ and $\gamma^1 \in \mathbf{R}^{4\times 4}$ contain unrestricted parameters.

Thus, the resulting unrestricted four-factor AFNS-R model has $\mathbb{P}$-dynamics given by the following:

$$\begin{pmatrix} dL_t \\ dS_t \\ dC_t \\ dX_t^R \end{pmatrix} = \begin{pmatrix} \kappa_{11}^{\mathbb{P}} & \kappa_{12}^{\mathbb{P}} & \kappa_{13}^{\mathbb{P}} & \kappa_{14}^{\mathbb{P}} \\ \kappa_{21}^{\mathbb{P}} & \kappa_{22}^{\mathbb{P}} & \kappa_{23}^{\mathbb{P}} & \kappa_{24}^{\mathbb{P}} \\ \kappa_{31}^{\mathbb{P}} & \kappa_{32}^{\mathbb{P}} & \kappa_{33}^{\mathbb{P}} & \kappa_{34}^{\mathbb{P}} \\ \kappa_{41}^{\mathbb{P}} & \kappa_{42}^{\mathbb{P}} & \kappa_{43}^{\mathbb{P}} & \kappa_{44}^{\mathbb{P}} \end{pmatrix} \left(\begin{pmatrix} \theta_1^{\mathbb{P}} \\ \theta_2^{\mathbb{P}} \\ \theta_3^{\mathbb{P}} \\ \theta_4^{\mathbb{P}} \end{pmatrix} - \begin{pmatrix} L_t \\ S_t \\ C_t \\ X_t^R \end{pmatrix} \right) dt + \Sigma \begin{pmatrix} dW_t^{L,\mathbb{P}} \\ dW_t^{S,\mathbb{P}} \\ dW_t^{C,\mathbb{P}} \\ dW_t^{R,\mathbb{P}} \end{pmatrix}.$$

This is the transition equation in the extended Kalman filter estimation.

4.2 Model Estimation and Econometric Identification

Due to the nonlinear relationship between state variables and bond prices in equation (4), the model cannot be estimated with the standard Kalman filter. Instead, we use the extended Kalman filter as in Kim and Singleton (2012); see Christensen and Rudebusch (2019) for details. Furthermore, to make the fitted errors comparable across bonds of various maturities, we scale each bond price by its duration. Thus, the measurement equation for the bond prices takes the following form:

$$\frac{P_t^i(t_0^i,\tau^i)}{D_t^i(t_0^i,\tau^i)} = \frac{\widehat{P}_t^i(t_0^i,\tau^i)}{D_t^i(t_0^i,\tau^i)} + \varepsilon_t^i.$$

Here, $\widehat{P}_t^i(t_0^i,\tau^i)$ is the model-implied price of bond i, $D_t^i(t_0^i,\tau^i)$ is its duration, which is calculated before estimation, and ε_t^i represents independent and Gaussian distributed measurement errors with mean zero and a common standard deviation σ_ε. See Andreasen et al. (2019) for evidence supporting this formulation of the measurement equation.

Furthermore, since the market-wide risk factor across the bond-specific premia is a latent factor that we do not observe, its level is not identified without additional restrictions. We therefore let the first thirty-year, 4 percent coupon Confederation bond, which was issued on April 8, 1998, and matures on April 8, 2028, have a unit loading on this factor. That is, $\beta^i = 1$ for this security. This choice implies that the β^i sensitivity parameters measure the sensitivity to this factor relative to that of the thirty-year 2028 Confederation bond.

Finally, we note that the $\lambda^{R,i}$ parameters can be hard to identify if their values are too large or too small. Consequently, we impose the restriction that they fall within the range from 0.0001 to 10, which is a restriction that has no practical consequences. In addition, for numerical stability during model optimization, we impose the restriction that the β^i parameters fall within the range from 0 to 250.

4.3 Estimation Results

This section presents our benchmark estimation results. In the interest of simplicity, we devote the paper's focus to a version of the AFNS-R model in which $K^{\mathbb{P}}$ and Σ are diagonal matrices.

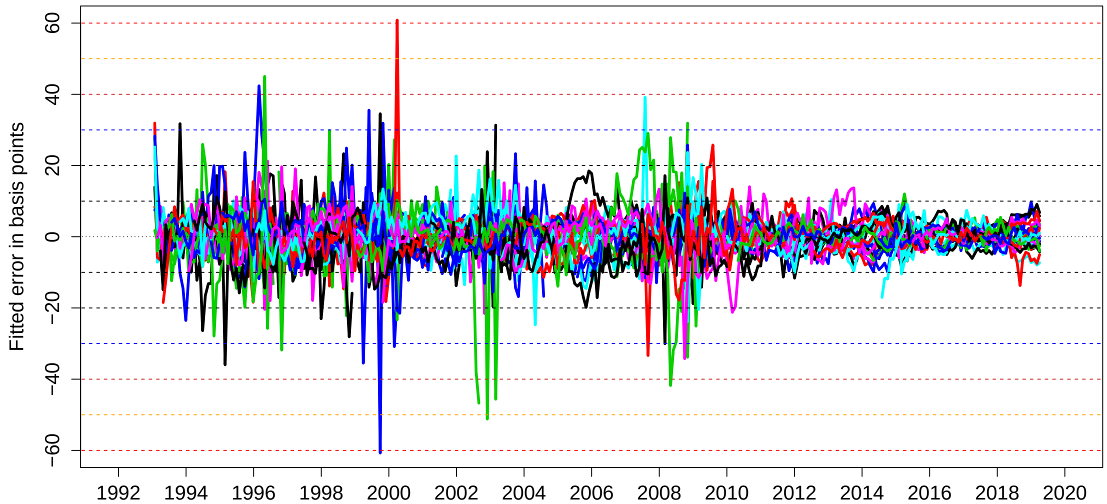


Figure 4: **Fitted Errors of Swiss Confederation Bond Yields**
Illustration of the fitted errors of Swiss Confederation bond yields-to-maturity implied by the AFNS-R model estimated with all bonds. The data are monthly and cover the period from January 29, 1993, to March 29, 2019.

As shown in ACR, these restrictions have hardly any effects on the estimated bond-specific risk premia, because they are identified from the model's $\mathbb{Q}$-dynamics, which are independent of $K^{\mathbb{P}}$ and display only a weak link to Σ through the small convexity adjustment in the yields. Furthermore, we confirm this later on in our robustness analysis in Section 5.2 when we consider more flexible specifications of the AFNS-R model's $\mathbb{P}$-dynamics.

The impact of accounting for the bond-specific risk premia is apparent in our results. To examine the model fit and to make pricing errors comparable accross securities, we compute the pricing errors based on the implied yield on each coupon bond. For the price on the ith coupon bond $P_t^i(\tau, C^i)$, we find the value of $y_t^{i,c}$ that solves the following:

$$P_t^i(\tau^i, C^i) = C^i(t_1 - t)\exp\left\{-y_t^{i,c}(t_1 - t)\right\} + \sum_{j=2}^{N} C^i \exp\left\{-y_t^{i,c}(t_j - t)\right\} + \exp\left\{-y_t^{i,c}(t_N - t)\right\}. \quad (5)$$

For the model-implied estimate of this bond price, denoted $\hat{P}_t^i(\tau, C^i)$, we find the corresponding implied yield $\hat{y}_t^{i,c}$ and report the pricing error as $y_t^{i,c} - \hat{y}_t^{i,c}$. The bond pricing errors produced by the AFNS model indicate a reasonable fit with an overall root mean-squared error (RMSE) of 9.08 basis points. However, we see a substantial improvement in the pricing errors when correcting for the bond-specific risk premia, as the AFNS-R model has a much lower overall RMSE of just 5.78 basis points.[17]

[17]The full details of the model fit comparison are provided in online Appendix B.

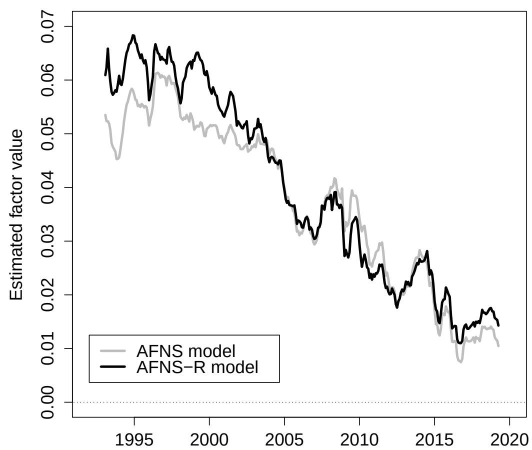
Parameter	AFNS		AFNS-R	
	Est.	SE	Est.	SE
$\kappa_{11}^{\mathbb{P}}$	0.0103	0.0286	0.0097	0.0373
$\kappa_{22}^{\mathbb{P}}$	0.2065	0.1234	0.1924	0.1352
$\kappa_{33}^{\mathbb{P}}$	0.8432	0.2537	0.5012	0.2880
$\kappa_{44}^{\mathbb{P}}$	-	-	0.0072	0.0345
σ_{11}	0.0028	0.0000	0.0033	0.0001
σ_{22}	0.0077	0.0002	0.0088	0.0006
σ_{33}	0.0162	0.0008	0.0191	0.0014
σ_{44}	-	-	0.0114	0.0010
$\theta_1^{\mathbb{P}}$	0.0345	0.0061	0.0423	0.0161
$\theta_2^{\mathbb{P}}$	-0.0242	0.0063	-0.0263	0.0105
$\theta_3^{\mathbb{P}}$	-0.0137	0.0039	0.0120	0.0122
$\theta_4^{\mathbb{P}}$	-	-	0.1232	0.2795
λ	0.2746	0.0026	0.1881	0.0032
$\kappa_R^{\mathbb{Q}}$	-	-	2.9361	0.1404
$\theta_R^{\mathbb{Q}}$	-	-	-0.0057	0.0004
σ_ε	0.0010	3.38×10^{-6}	0.0007	3.89×10^{-6}
Max $\mathcal{L}^{EKF}$	35,445.24		37,527.98	

Table 2: **Estimated Dynamic Parameters**
The table shows the estimated dynamic parameters for the AFNS and AFNS-R models estimated with a diagonal specification of $K^{\mathbb{P}}$ and Σ.

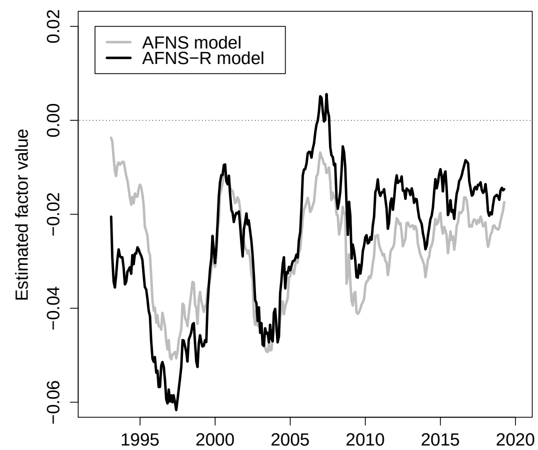
The time series of the individual fitted yield error series from the AFNS-R model are shown in Figure 4. Since 2010 the model has been able to deliver a very accurate fit to the entire cross section of bond yields, effectively keeping all bonds within a 10-basis-point error band. Thus, neither the 2009-2014 period with shorter-term interest rates constrained by the ZLB nor the negative rates prevailing since December 2014 appear to cause the model any problems.

Table 2 reports the estimated dynamic parameters. With the exception of the curvature factor, which is notably more persistent and has a higher mean in the AFNS-R model, the dynamics of the first three factors are qualitatively very similar across the two estimations. Furthermore, λ is significantly smaller in the AFNS-R model. This implies that the yield loadings of the slope factor decay toward zero at a slower pace as maturity increases. At the same time, the peak of the curvature yield loadings is located at a later maturity. Thus, in the AFNS-R model, slope and curvature matter more for longer-term yields. Since long-term yields tend to be more persistent than short-term yields, this helps explain the greater persistence of the curvature factor in the AFNS-R model.

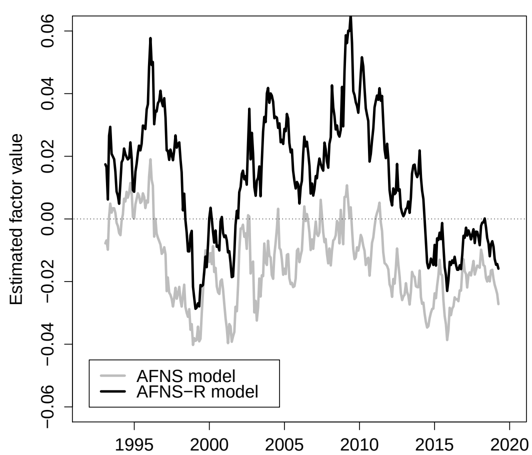
The estimated paths of the level, slope, and curvature factors from the two models are shown in Figure 5. The two models' level and slope factors are fairly close to each other during the entire sample, while there are notable level differences between the two estimated curvature factors throughout most of the sample period. Accordingly, the main impact of



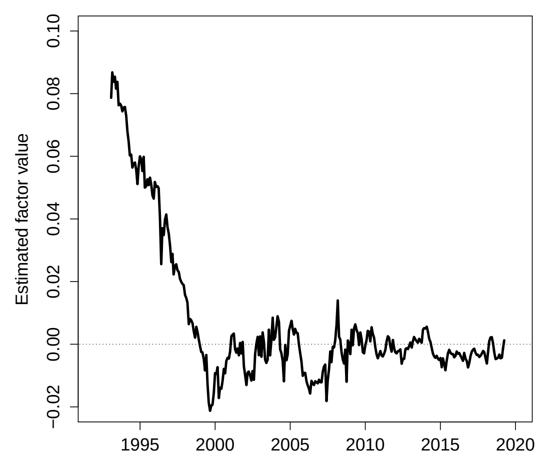
(a) L_t



(b) S_t



(c) C_t



(d) X_t^R

Figure 5: **Estimated State Variables**
Illustration of the estimated state variables from the AFNS and AFNS-R models.

accounting for the bond-specific risk premia is on the curvature of the frictionless yield curve. The marketwide factor driving the variation in the bond-specific risk premia unique to the AFNS-R model is very persistent with moderate volatility. Its estimated path exhibits a persistent decline until the late 1990s and has fluctuated close to zero since then.

4.4 Analysis of Model Fit

In this section, we aim to provide some perspective on how well the AFNS-R model fits the Swiss Confederation bond price data. Therefore, we compare its fit to that of a set of well-established yield curve models. Specifically, we compare it to the AFNS model already

Maturity bucket	No. obs.	AFNS		DNS		Svensson		AFGNS		AFNS-R	
		Mean	RMSE	Mean	RMSE	Mean	RMSE	Mean	RMSE	Mean	RMSE
0-2	784	-0.39	15.61	-0.45	14.78	-0.58	13.67	-0.03	10.34	0.08	9.56
2-4	924	0.15	9.77	0.80	9.46	0.83	7.92	0.03	7.97	0.21	6.06
4-6	935	-0.46	8.12	-0.81	7.89	-0.39	7.41	-0.68	7.28	-0.89	5.62
6-8	838	0.34	6.08	-0.41	6.16	-0.28	5.75	0.11	5.96	0.73	4.48
8-10	773	-0.36	6.09	-1.03	6.71	-0.87	4.82	-0.14	4.78	0.67	4.20
10-12	583	-0.30	6.74	-0.67	7.01	-0.64	5.61	0.05	5.38	0.45	4.03
12-14	334	-0.53	6.63	-0.12	6.64	0.47	4.46	0.58	4.37	0.03	3.72
14-16	212	1.24	7.62	2.30	7.68	2.08	5.41	2.00	5.98	0.23	4.96
16-18	181	1.56	6.75	3.07	6.40	2.05	4.36	1.31	4.51	0.00	4.96
18-20	214	2.72	7.90	4.50	8.17	2.75	6.76	2.29	6.95	1.05	5.55
20-22	120	-0.65	6.08	1.96	5.84	1.51	5.52	-0.11	5.59	0.74	5.08
22-24	132	-3.09	7.24	-0.57	7.20	-1.60	4.30	-3.38	5.36	-1.44	4.54
24-26	108	-3.70	8.51	-1.80	8.28	-1.50	6.46	-3.30	7.22	-0.13	5.80
26-28	118	-7.55	9.64	-4.68	8.41	-5.09	6.98	-7.38	8.91	-1.96	6.28
28-30	90	-5.33	10.98	-3.75	12.07	-6.80	8.20	-7.60	9.03	-2.24	5.56
30<	344	1.65	10.14	-1.01	10.77	-0.64	6.40	0.51	6.83	0.06	5.58
All bonds	6,690	-0.22	9.08	-0.18	8.95	-0.21	7.56	-0.22	7.03	0.09	5.78

Table 3: **Summary Statistics of Fitted Errors of Swiss Confederation Bond Yields**
This table reports the mean pricing errors (Mean) and the root mean-squared pricing errors (RMSE) of the Swiss bond prices for various models estimated from the sample of Swiss Confederation bond prices. The pricing errors are reported in basis points and computed as the difference between the implied yield on the coupon bond and the model-implied yield on this bond. The data are monthly and cover the period from January 29, 1993, to March 29, 2019.

considered, the dynamic Nelson-Siegel (DNS) model of Diebold and Li (2006), a similar dynamic version of the Svensson (1995) model, and the arbitrage-free generalized Nelson-Siegel (AFGNS) model described in Christensen et al. (2009).[18]

Table 3 evaluates the ability of all these models to match the market prices of the coupon bonds. As before, the pricing errors are computed based on the implied yield on each coupon bond to make these errors comparable across securities.

Table 3 reports the summary statistics for the fit to all bonds in the sample from the five model estimations broken down into maturity buckets. The AFNS-R model provides the closest fit to the data across the entire term structure. Compared to the AFNS and DNS models, which have only three factors, this result is not surprising. However, this result may be a little bit surprising compared to the flexible five-factor AFGNS model. Importantly, these results underscore a couple of takeaways. First, they show that it is crucial to account for bond-specific risk premia when it comes to modeling Swiss Confederation bond prices. Second, the really competitive fit of the AFNS-R model suggests that this model structure is well specified and strikes an appropriate balance between the number of frictionless factors and the bond-specific risk premium components. This view is supported further by our robustness analysis in Section 5.2.

[18] In online Appendix C, we also use these models to assess the quality of our Swiss data by comparing the model-implied yield curves to those produced by SNB staff and made publicly available on the SNB website.

5 The Confederation Bond Safety Premium

In this section, we analyze the Confederation bond safety premia implied by the estimated AFNS-R model described in the previous section. First, we formally define the bond safety premium, study its historical evolution, and explore its robustness to alternative modeling assumptions, including accounting for liquidity risk. We then use regression analysis to assess whether the safety premium experienced a structural break following the launch of the euro in January 1999 and the introduction of negative rates in Switzerland in December 2014, and whether it has been affected by recent ECB safe-asset purchases. We end the section with an analysis of the term structure of safety premia.

5.1 The Estimated Confederation Bond Safety Premium

We now use the estimated AFNS-R model to extract the safety premium in the Confederation bond market. To compute this premium, we first use the estimated parameters and the filtered states $\{X_{t|t}\}_{t=1}^{T}$ to calculate the fitted Confederation bond prices $\left\{\hat{P}_t^i\right\}_{t=1}^{T}$ for all outstanding securities in our sample. These bond prices are then converted into yields to maturity $\left\{\hat{y}_t^{c,i}\right\}_{t=1}^{T}$ by solving the following fixed-point problem:

$$\hat{P}_{t=1}^i = C(t_1 - t)\exp\left\{-(t_1 - t)\hat{y}_t^{c,i}\right\} + \sum_{k=2}^{n} C\exp\left\{-(t_k - t)\hat{y}_t^{c,i}\right\} + \exp\left\{-(T - t)\hat{y}_t^{c,i}\right\}, \quad (6)$$

for $i = 1, 2, ..., n$, which entails that $\left\{\hat{y}_t^{c,i}\right\}_{t=1}^{T}$ is approximately the rate of return on the ith Confederation bond if held until maturity (see Sack and Elsasser 2004). To obtain the corresponding yields without correcting for the safety premium, a new set of model-implied bond prices are computed from the estimated AFNS-R model but using only its frictionless part, i.e., with the constraints that $X_{t|t}^R = 0$ for all t, $\sigma_{44} = 0$, and $\theta_R^{\mathbb{Q}} = 0$. These prices are denoted $\left\{\tilde{P}_t^i\right\}_{t=1}^{T}$ and converted into yields to maturity $\tilde{y}_t^{c,i}$ by using (6). They represent estimates of the prices that would prevail in a world without any financial frictions or convenience yields. The safety premium for the ith Confederation bond is then defined as follows:

$$\Psi_t^i \equiv \tilde{y}_t^{c,i} - \hat{y}_t^{c,i}. \quad (7)$$

Figure 6 shows the average Swiss Confederation bond safety premium $\bar{\Psi}_t$ across the outstanding Confederation bonds at each point in time. The average estimated safety premium clearly varies notably over time with a maximum of around 100 basis points achieved in the summer of 2007 at the onset of the financial crisis and again following the bankruptcy of Lehman Brothers in September 2008.

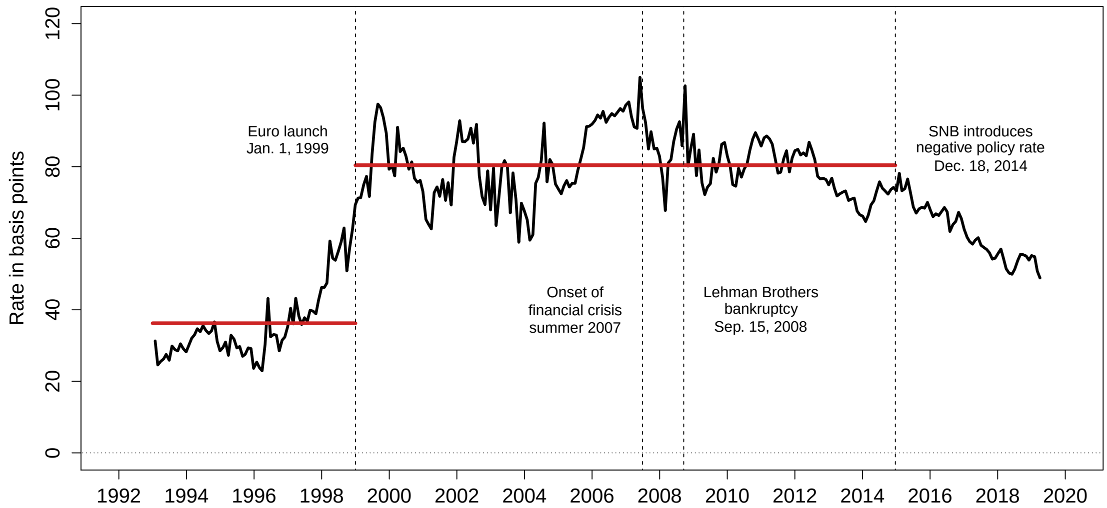


Figure 6: **Average Estimated Swiss Confederation Bond Safety Premium**
Illustration of the average estimated Swiss Confederation bond safety premium for each observation date, as implied by the AFNS-R model. The Confederation bond safety premia are measured as the estimated yield difference between the frictionless yield to maturity of individual Confederation bonds with the market risk factor turned off and the corresponding fitted yield to maturity. The data cover the period from January 29, 1993, to March 31, 2019. The solid red lines indicate the average of the safety premium series from January 31, 1993, to December 31, 1998, and from January 31, 1999, to November 30, 2014.

Interestingly, however, there seems to be a persistent upward shift in the safety premium happening somewhere in the period between late 1997 and mid-1999. Such a dramatic shift could be linked to the introduction of the euro, which may have reduced the safe-haven attractiveness of euro-area assets by possibly changing perceptions about implicit risk-sharing between core and noncore eurozone countries. Consequently, the introduction of the euro might have increased the appeal of Swiss assets, including Swiss Confederation bonds, as safe-haven assets.

In addition, note that the average safety premium appears to have trended lower ever since the SNB introduced negative interest rates. In an effort to curb the appreciation of the Swiss franc by making investments in Swiss assets less attractive, the SNB lowered the interest rate on sight deposits, first to -25 basis points in December 2014, and then to -75 basis points in January 2015.[19] In response, the entire yield curve of Swiss Confederation bonds shifted downwards, as documented in Christensen (2019), which increased the opportunity costs of holding these bonds and might have reduced their safety premium.

Since safe asset purchases by foreign central banks tend to reduce the available stock of

[19]See SNB's press releases from December 18, 2014, and January 15, 2015, respectively.

foreign safe assets, the government bond purchases made by the ECB since January 2015 through its public sector purchase program (PSPP) could have made Swiss Confederation bonds relatively less attractive and reduced the Swiss safety premium. Alternatively, these asset purchases could have reduced the available stock of safe assets globally and hence could have raised the premium of all safe assets. Which of these two effects is more important is an empirical question we aim to explore as well, but first we examine the robustness of the estimated safety premium series.

5.2 Robustness Analysis

This section analyzes the robustness of the previously reported average safety premium to some of the main assumptions imposed so far. Throughout this section, the AFNS-R model with diagonal $K^{\mathbb{P}}$ and Σ matrices serves as the benchmark.

5.2.1 Flexible P-Dynamics, Different Data Frequency and Stochastic Volatility

First, we assess whether the specification of the factor dynamics within the AFNS-R model matters for the estimated Swiss Confederation bond safety premium. To do so, we estimate the AFNS-R model with unrestricted $K^{\mathbb{P}}$ and diagonal Σ matrix and the AFNS-R model with unconstrained dynamics, i.e., the AFNS-R model with unrestricted $K^{\mathbb{P}}$ and lower triangular Σ matrix.

Figure 7 shows the estimated Swiss safety premium series from these two estimations and compares them to the series produced by our benchmark specification. Note that the three series are barely distinguishable. This shows that the specification of the $\mathbb{P}$-dynamics within the AFNS-R model plays only a very modest role for the estimated bond-specific risk premia, which is consistent with the findings of ACR in the context of U.S. TIPS. This result also provides support for our choice to focus only on the most parsimonious specification of the AFNS-R model in our empirical analysis.

Second, we assess whether the data frequency plays any role for our results by estimating the AFNS-R model using daily, weekly, and monthly data. Building on the results above, we focus on the most parsimonious AFNS-R model with diagonal $K^{\mathbb{P}}$ and Σ matrices.

Figure 8 shows the estimated Swiss safety premium series from all three estimations. Note that they are hardly distinguishable. Hence, the data frequency matters little for our results. Clearly, at the higher daily and weekly frequencies, there are a few isolated spikes that are absent in the monthly series, but they are too infrequent to have an impact on the estimation results.

Finally, we explore whether allowing for stochastic volatility in one or more of the fundamental frictionless factors within the AFNS-R model affects the estimated Swiss Confederation bond safety premium. Specifically, we consider the four admissible combinations of allowing for spanned stochastic volatility generated by one or two factors in the model follow-

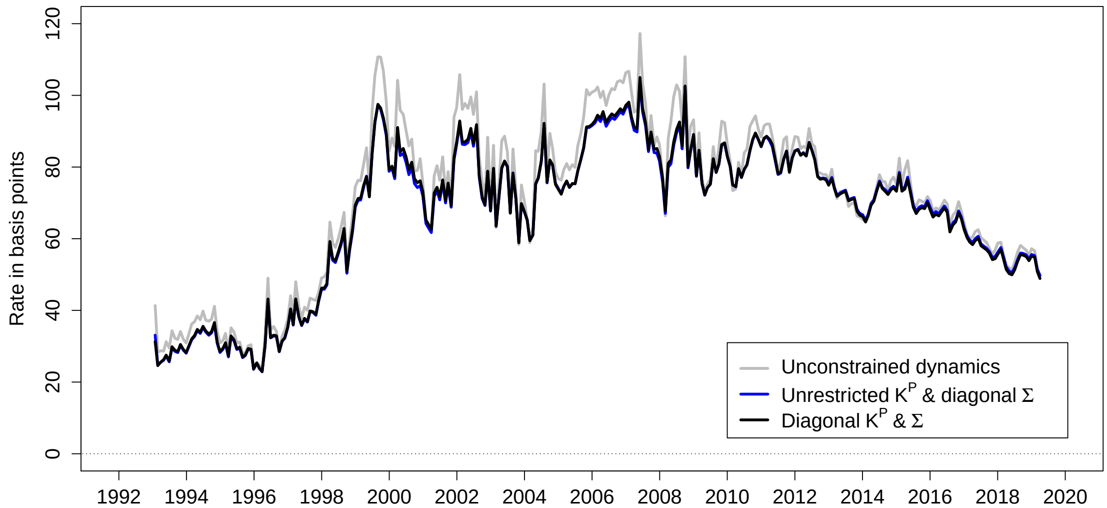


Figure 7: **Average Estimated Swiss Confederation Bond Safety Premium: Alternative $\mathbb{P}$ Dynamics**

Illustration of the average estimated Swiss Confederation bond safety premium for each observation date, as implied by the AFNS-R model when estimated with three specifications of its dynamics as detailed in the text. In all cases the Confederation bond safety premia are measured as the estimated yield difference between the fitted yield to maturity of individual Confederation bonds and the corresponding frictionless yield to maturity with the liquidity risk factor turned off. The data cover the period from January 29, 1993, to March 29, 2019.

ing Christensen et al. (2014). In light of the results above, we focus on the most parsimonious specification of each model with diagonal $K^{\mathbb{P}}$ and Σ matrices.

We refer to these models as AFNS models because they share the key properties of the AFNS model. First, the three frictionless state variables have joint dynamics under the risk-neutral probability measure used for pricing closely matching the arbitrage-free Nelson-Siegel models described in Christensen et al. (2011). Furthermore, the frictionless short rate remains defined as $r_t = L_t + S_t$. Therefore, to keep the notation simple, we use AFNS(i) to denote an AFNS model as defined above, where i refers to the number of factors generating stochastic volatility, while the letters—L, S, and C—are used to indicate the source(s) of stochastic volatility in the model.

Figure 9 shows the estimated Swiss safety premium series from these estimations. Note again that they are barely distinguishable. Thus, allowing for stochastic volatility within the AFNS-R model only plays a very modest role for our results. This provides support for our choice to focus only on the Gaussian AFNS-R model with constant volatility.

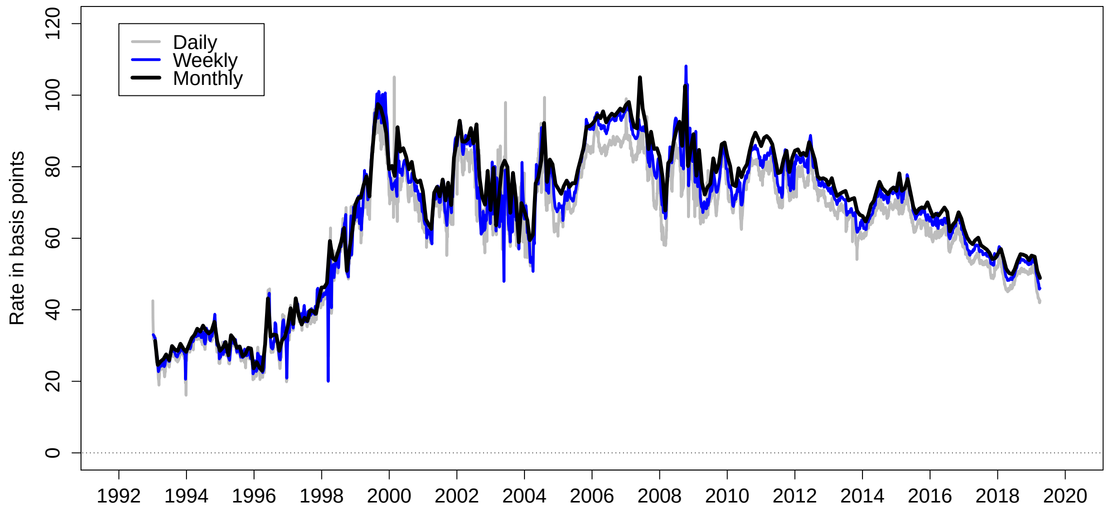


Figure 8: **Average Estimated Swiss Confederation Bond Safety Premium: Data Frequency**
Illustration of the average estimated Swiss Confederation bond safety premium for each observation date implied by the AFNS-R model when estimated using daily, weekly, and monthly data. In all cases the Confederation bond safety premia are measured as the estimated yield difference between the fitted yield to maturity of individual Confederation bonds and the corresponding frictionless yield to maturity with the liquidity risk factor turned off. The data cover the period from January 4, 1993, to March 29, 2019.

5.2.2 Accounting for Liquidity Risk

Despite the robustness of the AFNS-R model results documented so far, we acknowledge that one could be concerned about the role of liquidity risk for our findings given the latent-factor nature of the model. Therefore, to address this important question, we estimate an expanded version of the AFNS-R model that includes a factor intended to capture the liquidity risk of the Swiss Confederation bonds. Specifically, we follow D'Amico et al. (2018) and model their liquidity risk using a reduced-form approach with a liquidity-adjusted short rate given by

$$\overline{r}_t^i = r_t + \beta^i (1 - e^{-\lambda^{R,i}(t - t_0^i)}) X_t^R + X_t^{liq},$$

where the liquidity risk factor X_t^{liq} is an independent Ornstein-Uhlenbeck process under the risk-neutral $\mathbb{Q}$-measure

$$dX_t^{liq} = \kappa_{liq}^{\mathbb{Q}} (\theta_{liq}^{\mathbb{Q}} - X_t^{liq}) dt + \sigma_{55} dW_t^{liq,\mathbb{Q}},$$

while the objective $\mathbb{P}$-dynamics are obtained in the standard way as before, see the appendix for details. We refer to this expanded five-factor model as the AFNS-R-L model.

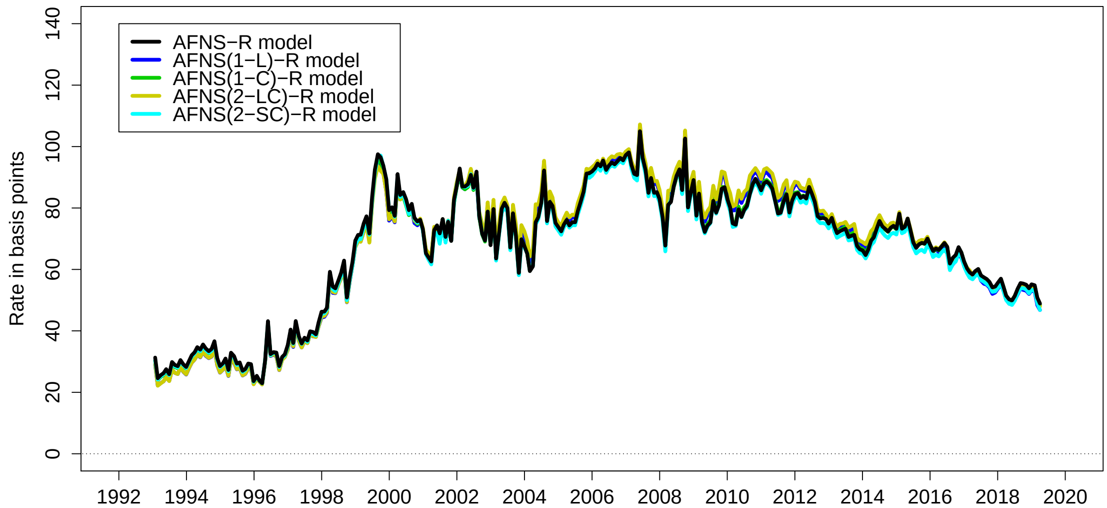


Figure 9: **Average Estimated Swiss Confederation Bond Safety Premium: Allowing for Stochastic Volatility**

Illustration of the average estimated Swiss Confederation bond safety premium for each observation date as implied by the AFNS-R model when estimated with and without allowing for stochastic volatility as detailed in the text. In all cases the Confederation bond safety premia are measured as the estimated yield difference between the fitted yield to maturity of individual Confederation bonds and the corresponding frictionless yield to maturity with the liquidity risk factor turned off. The data cover the period from January 29, 1993, to March 29, 2019.

When we estimate this fully latent-factor model, it turns out that it struggles to distinguish between liquidity and safety premia as we document below. Therefore, we also estimate a version of the model where X_t^{liq} is assumed to be observable, but with error.

Our observable composite liquidity factor has three components. First, inspired by the analysis of Hu et al. (2013), we include a noise measure of the Swiss Confederation bond prices to capture variation in the amount of arbitrage capital available in the market for these bonds. Second, we use a standard measure of realized yield volatility to proxy for bond liquidity as recommended by Houweling et al. (2005). As the final proxy for bond liquidity, we use bid and ask prices of the Swiss Confederation bonds.

To construct our composite liquidity factor, we follow the approach of Abrahams et al. (2016) and standardize all three variables before calculating the average. To smooth the series further and make it closer to the yield data used in the model estimation, we calculate the four-week moving average. Finally, to align the data with our monthly yield data, we focus on the value at the end of each month and deduct the minimum value of this series to eliminate negative values. We define the resulting series shown in Figure 10 as our monthly observable composite liquidity factor, see the appendix for all details.

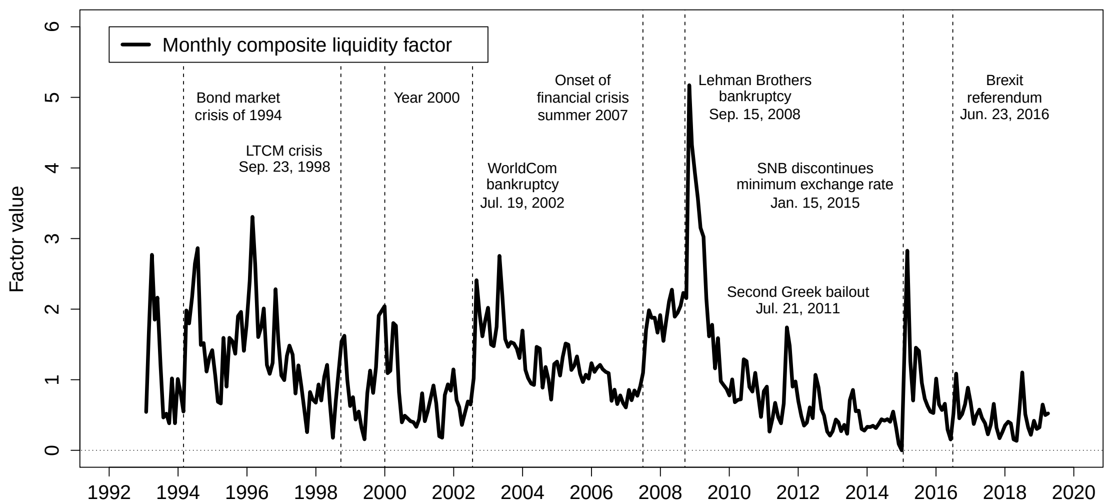


Figure 10: **Monthly Observable Composite Liquidity Factor**

We note that the composite liquidity factor is characterized by notable spikes around a number of well-documented financial market liquidity shocks, in particular the 1994 bond market crash, the 1998 LTCM crisis, the WorldCom and Lehman Brothers bankruptcies, and the SNB's abandonment of its minimum exchange rate to the euro in January 2015. As a consequence, we feel comfortable relying on this observable composite factor to account for the liquidity risk of the Confederation bonds within the AFNS-R-L model.

To align the observable liquidity factor with the yield data, we add a scaling parameter δ^L in the formulation of the liquidity-adjusted short rate[20]

$$\overline{r}_t^i = r_t + \beta^i(1 - e^{-\lambda^{R,i}(t-t_0^i)})X_t^R + \delta^L X_t^{liq}.$$

Furthermore, it turns out to be critical to restrict the mean-reversion rate of X_t^{liq} under the risk-neutral $\mathbb{Q}$-measure, $\kappa_{liq}^{\mathbb{Q}}$. Without any restrictions on this parameter, the likelihood optimizer assigns it unreasonably large values of 20 or more, which implies that the estimated liquidity premia are essentially constant over time. The interpretation of this outcome is that the variation in the liquidity factor is hard to reconcile with the variation in the bond price data, and a really high value of $\kappa_{liq}^{\mathbb{Q}}$ is a way for the optimizer to avoid dealing with the variation in X_t^{liq}. Furthermore, in that case, it can easily offset any effects from the X_t^{liq}-component in the fitted bond prices by an opposite movement in the level factor L_t. Therefore, to force the optimizer to meaningfully engage with the variation of X_t^{liq}, we impose the restriction that its mean-reversion is the same under both probability measures, $\kappa_{55}^{\mathbb{P}} = \kappa_{liq}^{\mathbb{Q}}$.

[20]The scaling parameter δ^L is not needed when X_t^{liq} is treated as a latent factor and therefore fixed at 1.

Finally, to allow the optimizer to further reconcile the variation in X_t^{liq} with that of the bond prices, we assume that X_t^{liq} is observed with error. Thus, we add the measurement equation

$$X_t^{liq} = \widehat{X}_{t|t-1}^{liq} + \varepsilon_t^{liq},$$

where ε_t^{liq} is assumed to be normally distributed with a mean equal to zero and a standard deviation of $\sigma_{\varepsilon^{liq}}$, which we fix at 0.35, or about one half of the standard deviation of X_t^{liq}.[21] This choice leaves the optimizer some room to deviate from the observed values of X_t^{liq}, while still being forced to meaningfully engage with its variation and try to reconcile it with the variation in the bond prices.

To compute the estimated liquidity premia implied by the AFNS-R-L model, we first use the estimated parameters and the filtered states $\left\{X_{t|t}\right\}_{t=1}^{T}$ to calculate the frictionless bond prices $\left\{\tilde{P}_t^i\right\}_{t=1}^{T}$ for all outstanding securities in our sample, i.e., we are just relying on the estimated dynamics of (L_t, S_t, C_t). These bond prices are then converted into yields to maturity $\left\{\tilde{y}_t^{c,i}\right\}_{t=1}^{T}$ by solving the fixed-point problem in equation (6) for $i = 1, 2, ..., n$. To obtain the corresponding yields with the liquidity premium, a new set of model-implied bond prices are computed from the estimated AFNS-R-L model using the liquidity-adjusted discount function

$$\overline{r}_t = r_t + \delta^L X_t^{liq},$$

where $r_t = L_t + S_t$ remains the frictionless short rate as before.[22] These prices are denoted $\left\{\overline{P}_t^i\right\}_{t=1}^{T}$ and converted into yields to maturity $\overline{y}_t^{c,i}$ using equation (6). The liquidity premium for the ith bond is then defined as

$$\Lambda_t^i \equiv \overline{y}_t^{c,i} - \tilde{y}_t^{c,i}.$$

Figure 11 shows the average Swiss Confederation bond liquidity premium $\bar{\Lambda}_t$ across the outstanding bonds at each point in time from four different implementations of the AFNS-R-L model. In the first implementation, X_t^{liq} is assumed to be a latent factor, while in the three other implementations X_t^{liq} is assumed observable, but with error, and with three different values for the scaling parameter $\delta^L \in \{0.005, 0.0075, 0.01\}$.

Two important findings stand out from this figure. First, the average estimated liquidity premium series from the AFNS-R-L model with all factors treated as latent is very different from the series obtained with X_t^{liq} treated as observable, but with error. Most notably, the former series *declines* following the Lehman Brothers bankruptcy. We take this and other unusual spikes as evidence that this particular model implementation struggles to distinguish between liquidity and safety premium components in the bond prices. Still, we consider its estimated average liquidity premium close to 50 basis points to be a realistic value. For

[21] The standard deviation of the monthly X_t^{liq} series is 0.7298.

[22] This approach mimics how we calculate the safety premia within the AFNS-R model.

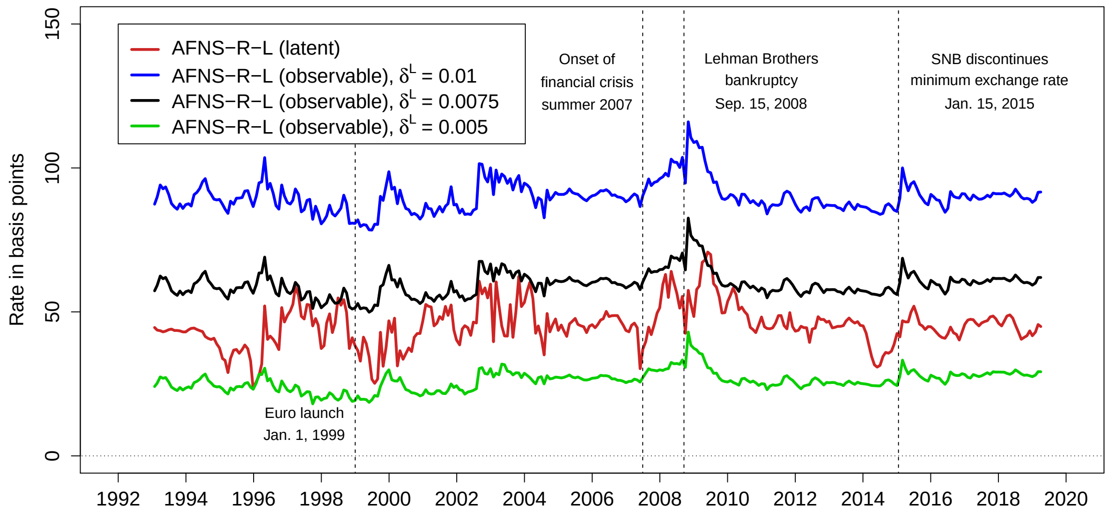


Figure 11: **Average Estimated Swiss Confederation Bond Liquidity Premium**
Illustration of the average estimated Swiss Confederation bond liquidity premium for each observation date implied by the AFNS-R-L model estimated with X_t^{liq} as a latent factor and three versions with X_t^{liq} assumed observable, but with error. The Confederation bond liquidity premia are measured as the estimated yield difference between the frictionless yield to maturity of individual Confederation bonds with both the bond-specific and liquidity risk factors turned off and the corresponding fitted yield to maturity with only the bond-specific risk factor turned off. The data cover the period from January 31, 1993, to March 31, 2019.

comparison, ACR report an average liquidity premium of 34 basis points for the U.S. TIPS market.

The other important finding is that the scaling parameter δ^L matters for the size of the estimated liquidity premia when X_t^{liq} is treated as observable, but with error, as one could expect. Ideally, we would like to leave δ^L as a free parameter to be estimated by the data, but unfortunately the optimizer trades off increases in δ^L with lower values of the level factor L_t to obtain higher likelihood values, but with absurdly high values of the estimated liquidity premia. Crucially, this matters little for the estimated safety premia as we document below. As a consequence, we treat δ^L as a fixed parameter and choose 0.0075 as our benchmark value since this choice produces an average liquidity premium close to that from the latent-factor implementation of the AFNS-R-L model as shown in Figure 11.

To compute the safety premium in the augmented AFNS-R-L model, we first use its estimated parameters and filtered states $\{X_{t|t}\}_{t=1}^T$ to calculate the fitted bond prices $\left\{\hat{P}_t^i\right\}_{t=1}^T$ for all outstanding securities in our sample. These bond prices are then converted into yields to maturity $\left\{\hat{y}_t^{c,i}\right\}_{t=1}^T$ by solving the fixed-point problem in equation (6) for $i = 1, 2, ..., n$. To obtain the corresponding yields adjusted for the liquidity premium, a new set of model-

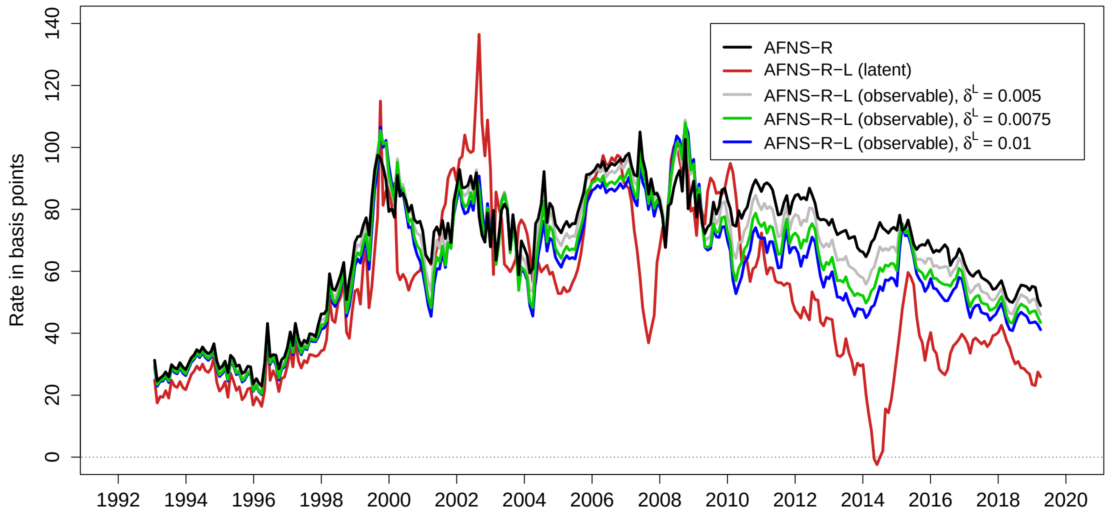


Figure 12: **Average Estimated Swiss Confederation Bond Safety Premium**
Illustration of the average estimated Swiss Confederation bond safety premium for each observation date implied by the AFNS-R and AFNS-R-L models. The Confederation bond safety premia are measured as the estimated yield difference between the frictionless yield-to-maturity of individual government bonds with both the bond-specific and liquidity risk factors turned off and the corresponding fitted yield-to-maturity with only the liquidity risk factor turned off. The data cover the period from January 31, 1993, to March 31, 2019.

implied bond prices are computed from the estimated AFNS-R-L model with the constraints that $X_{t|t}^{liq} = 0$ for all t as well as $\theta_{liq}^{\mathbb{Q}} = 0$ and $\sigma_{55} = 0$. These prices are denoted $\{\underline{P}_t^i\}_{t=1}^T$ and converted into yields to maturity $\underline{y}_t^{c,i}$ using equation (6).

The safety premium for the ith bond is then defined as

$$\Psi_t^i \equiv \hat{y}_t^{c,i} - \underline{y}_t^{c,i}. \tag{8}$$

Figure 12 shows the average Swiss Confederation bond safety premium $\bar{\Psi}_t$ across the outstanding bonds at each point in time from the same four AFNS-R-L model implementations already described earlier along with a comparison to the estimated safety premium series obtained with our baseline AFNS-R model.

First and most importantly, our baseline safety premium series from the AFNS-R model is very close to, and highly positively correlated with, the series produced by the AFNS-R-L model with X_t^{liq} treated as observable, but with error, and this conclusion is robust to varying the scaling parameter δ^L in the neighborhood of our benchmark value of 0.0075.

Second, the estimated safety premium series from the AFNS-R-L model with all factors treated as latent stands out as unusual both by its higher volatility and by its spikes up and down, reaching a maximum value in 2002 and low values in late 2007 and again in the spring of 2014. We take this as further evidence that this particular implementation of the AFNS-R-L model faces real challenges in distinguishing between liquidity and safety premia in the bond prices. This reinforces the impression that neither its liquidity premium series nor its safety premium series is convincing.

To summarize, we find the estimated safety premium series from the AFNS-R model to be robust to both alternative model dynamics and dynamic assumptions. Furthermore, it has modest sensitivity to the inclusion of an observable liquidity risk factor. As a consequence, we are comfortable using the parsimonious AFNS-R model with diagonal $K^{\mathbb{P}}$ and Σ matrices as our benchmark for the estimation of safety premia embedded in Confederation bond prices. Finally, given the persistence of the monetary policy changes we are interested in, we choose to focus on the results from monthly data, which greatly reduces the computational burden and simplifies the regression analysis detailed in the following section.

5.3 Regression Analysis

To test the hypotheses laid out in Section 5.1, we first look for structural breaks in the time series of the average Swiss safety premium. We run the CUSUM test from Page (1954) to measure the stability of the intercept coefficient in a regression of a vector of ones on the safety premium $\bar{\Psi}_t$. The test shows that there is one significant break in the premium series starting from March 1999 and lasting until the end of the sample. Therefore, the launch of the euro in January 1999 appears to have given rise to a break in the time series of the Swiss safety premium, while this is not the case for either the introduction of negative interest rates by the SNB or the ECB's asset purchases.

To explore this result further, we measure the average treatment effect of the launch of the euro, the introduction of negative interest rates, and the ECB's operation of the PSPP on Swiss safety premia within a regression framework. In particular, we run the following regression:

$$\bar{\Psi}_t = \alpha + \delta_{euro} d_t^{euro} + \delta_{nr} d_t^{nr} + \delta_{pspp} d_t^{pspp} + \delta_c D_t + \sum_{l=0}^{L} \delta_l X_{t-l} + \varepsilon_t, \tag{9}$$

where d_t^{euro} and d_t^{nr} are dummy variables that take the value of one onwards from January 1999 and from December 2014, respectively, and take the value of zero before then, while d_t^{pspp} is the stock of bonds acquired by the ECB through the PSPP, expressed as a percent of nominal GDP in the euro area;[23,24] D_t and X_t are vectors of a control dummy and of continuous

[23]We use nominal GDP for each calendar year except for 2019, for which we use a four-quarter rolling sum.

[24]In principle, this measure should only include the truly safe assets acquired by the ECB and should exclude bonds issued by high-risk countries in the periphery of the euro area. However, in the absence of that granular data, we use the entire stock of purchased government-backed securities as a proxy for the amount of absorbed safe assets. Provided the ratio of truly safe assets in the ECB's portfolio is close to being constant, which is a reasonable assumption given the fixed distribution key for the purchases, there should be little bias in the

variables, respectively. L is the number of lags included; and ε_t is a random residual. We consider the estimates of δ_{euro} and δ_{nr} to be the average treatment effects of the euro launch and the introduction of negative rates on the safety premium under the assumption that the confounding variables in the vector X_t are exogenous and that $E[\varepsilon_t|X_t] = 0$, while δ_{pspp} measures the effect on the safety premium of a 1 percentage point change in the stock of bonds held by the ECB.

We control for a host of confounding factors. In a core set of controls, we consider the CBOE Volatility Index (VIX), the spread between the Italian and the German 10-year government bond yield, the TED spread, and the on-the-run premium in U.S. Treasuries to proxy for investors' risk aversion, financial market uncertainty, and related demand for safe-haven assets;[25] the spread between the German and Swiss 10-year government bond yield proxies for the opportunity cost of holding Swiss Confederation bonds and the debt-to-GDP ratio is used to control for effects tied to the supply of Confederation bonds;[26] and the three-month CHF LIBOR is considered to proxy for the opportunity cost of holding money and the associated liquidity premiums of Confederation bonds, as explained in Nagel (2016). Furthermore, we include the average Confederation bond age and the one-month realized volatility of the ten-year Confederation bond yield as additional proxies for bond liquidity following the work of Houweling et al. (2005). Inspired by the analysis of Hu et al. (2013), we also include a noise measure of Swiss Confederation bond prices to control for the variation in the amount of arbitrage capital available in this market. Finally, we also add a dummy D_t that takes the value of one in the period of minimum exchange rate from September 2011 to January 2015 and zero otherwise, to control for any indirect effects of that particular SNB policy on the safety premium.

In addition to the set of core control variables, we consider several other confounding factors in the regressions. We add the overnight federal funds rate to proxy for the U.S. safe-asset liquidity premium as in Nagel (2016), and we use reported earnings per share of companies in the S&P 500 to account for opportunity costs in the equity market. We also consider the MOVE volatility index to proxy for risk aversion in the bond market. Finally, we include the total sight deposits at the SNB to control for any possible reserve-induced effects of the SNB's FX interventions; see Christensen and Krogstrup (2019).

Table 4 reports the results in which column (1) contains the outcomes without any controls, while columns (2) and (3) report the regressions conducted by using the core and extended group of controls, respectively.

The main finding is that the introduction of the euro had a significant *positive* effect on the Swiss safety premium. The estimated effect ranges from 25 to 44 basis points depending on the control variables included in the regression. On the other hand, the introduction of negative

estimated parameter.

[25]See Grisse and Nitschka (2015).

[26]See Nagel (2016) and Krishnamurthy and Vissing-Jorgensen (2012), respectively.

	1	2	3
α	36.2	113.9	279.6
	(0.00)	(0.00)	(0.00)
δ_{euro}	44.2	24.6	26.6
	(0.00)	(0.00)	(0.00)
δ_{nr}	-6.4	-17.6	-5.7
	(0.00)	(0.00)	(0.38)
δ_{pspp}	-1.1	-1.2	-1.6
	(0.00)	(0.00)	(0.00)
controls	no	core	all
Adj. R^2	0.82	0.89	0.93
DW	0.39	0.74	1.11

Table 4: **Average Treatment Effects of Euro Introduction and SNB's Negative Rates**

The table reports the coefficient estimates from regression (9) together with their respective p-values (in brackets) obtained using Newey-West standard errors. The first column reports the regression without controls, the second column reports the estimates with the core set of control variables, and the third column contains the results including all the controls. The last two rows report the adjusted R^2 and the Durbin-Watson statistic. The number of lags L in regressions (2) and (3) is set to 12. The sample starts in January 1993 and ends in March 2019.

interest rates by the SNB had a significant *negative* effect on the premium, but this result is not robust to including the extended set of control variables, most notably the sight deposits at the SNB, which grew substantially with the sizable amount of SNB interventions in the foreign exchange markets following the financial crisis. Finally, we do observe systematically significant *negative* effects from including the ECB's safe asset purchases under the PSPP in the regressions. The results suggest that an increase of 1 percentage point in these holdings lowers the Swiss safety premium by slightly more than 1 basis point. Hence, the cumulative effect between January 2015 and March 2019 lowered the Swiss safety premium by about 20 basis points.

We note that adding controls and their lags increases the adjusted R^2. Relatively high R^2s indicate that most confounding factors are controlled for and that our estimates of δ_{euro} could be largely considered as causal. However, if there are omitted variables correlated with the euro dummy, our estimate of the treatment effect would be biased upward. We therefore try to sharpen the inference regarding our three specific questions in the following sections.

5.3.1 Effects of the Introduction of the Euro in Isolation

To focus narrowly on the impact of the launch of the euro in isolation and avoid potentially polluting effects from the financial crisis, the Swiss introduction of negative rates, and the

Sample	No controls	Core controls	All controls
+/- 1 year	26.6	6.7	-53.1
	(0.00)	(0.65)	(0.11)
+/- 2 years	33.6	24.6	22.2
	(0.00)	(0.01)	(0.03)
+/- 3 years	36.3	30.6	36.4
	(0.00)	(0.00)	(0.00)
+/- 4 years	40.6	36.6	42.0
	(0.00)	(0.00)	(0.00)
+/- 5 years	40.3	33.7	41.3
	(0.00)	(0.00)	(0.00)
+/- 6 years	41.4	34.0	36.1
	(0.00)	(0.00)	(0.00)

Table 5: **Average Treatment Effects of the Euro Introduction in Isolation**
The table reports the estimated coefficients of the euro dummy variable in regression (10) together with their respective *p*-values (in brackets) obtained using Newey-West standard errors. The first column reports the regression without controls, the second column reports the estimates with the core set of control variables, and the third column contains the results including all the controls. The six samples start in January 1998, January 1997, January 1996, January 1995, January 1994, and January 1993, and end in December 1999, December 2000, December 2001, December 2002, December 2003, and December 2004, respectively.

ECB asset purchases, we run simplified regressions of the following form:[27]

$$\bar{\Psi}_t = \alpha + \delta_{euro} d_t^{euro} + \delta_l X_t + \varepsilon_t, \tag{10}$$

where we first use a sample that contains one year of data before the January 1, 1999, euro launch and one year of data after this date. We then expand this sample in a symmetric fashion by adding one more year of data before and after the euro launch and run the regression. We repeat this exercise until the sample contains six years of data before and after the euro launch, which is the maximum period given the January 1993 start date for our data. Overall, this allows us to compare the residual variation in the immediate pre-euro period to that during the immediate post-launch period.

The results are reported in Table 5. For the smallest (+/- 1 year) sample, the results are erratic, which may be due to low statistical power. However, for the most informative samples, defined as those containing at least three years of data before and after the euro launch, the estimated dummy coefficients are relatively stable, systematically highly statistically significant, and not sensitive to the set of controls used. Therefore, we see clear evidence of a statistically significant *positive* effect of roughly 35-40 basis points on the Swiss safety premium following the introduction of the euro.

[27] In these regressions, we have to drop the lags of the explanatory variables due to the limited number of observations and high number of control variables.

One might view this result as surprising given that investors knew the launch date well in advance and therefore could have rebalanced their portfolios towards Swiss assets earlier. However, there was lingering uncertainty about the rollout and the general implementation of the euro project even relatively close to the launch date. Furthermore, this happened in the context of general financial market uncertainty surrounding the Asian financial crisis of 1997 and the collapse of Long-Term Capital Management (LTCM) during the Russian sovereign debt crisis in the fall of 1998, both of which provided temporary boosts to the Swiss safety premium. Thus, it is not clear at what point and how fast we should expect to see a sustained fundamental change in the safety premium. To see no change would be an extreme example of irrational investor behavior that is not supported by our safety premium series as it had clearly started its upward trend before January 1, 1999. On the other hand, to see the full difference priced in ahead of the euro launch would require a level of foresight and rationality among market participants rarely documented in the empirical finance literature. Our results suggest that the effect starts to be reflected in the safety premium a couple of years before the launch and is fully priced in a couple of years after it. This seems reasonable given that the involved portfolio flows likely were informed by the general experience with the euro and perceptions about the operation of the new common monetary policy in the euro area.

5.3.2 Effects of the Introduction of Negative Rates in Isolation

To assess the impact of the effect in isolation of the introduction of negative rates, we follow a similar approach and run simplified regressions of the form:

$$\bar{\Psi}_t = \alpha + \delta_{nr} d_t^{nr} + \delta_l X_t + \varepsilon_t, \tag{11}$$

where we first use a sample that contains one year of data before the December 18, 2014, introduction of negative rates and one year of data after this date. As before, the regressions are repeated with symmetrically expanded samples that each time contain one more year of data before and after the December 18, 2014, event date. Given that our sample ends in March 2019, we can add up to four years of data around both sides of the event date in this exercise.

The results are reported in Table 6, in which we note a small and mostly insignificant effect on the Swiss safety premium following the introduction of negative rates in December 2014. Thus, any negative effects on the safety premium in recent years appear to be due to other factors.

5.3.3 Effects of the ECB's PSPP in Isolation

One particular factor we are interested in exploring further is the effect on the Swiss safety premium from the ECB's government bond purchases, which were announced in January

Sample	No controls	Core controls	All controls
+/- 1 year	0.5	-0.6	-4.6
	(0.83)	(0.89)	(0.29)
+/- 2 years	-2.7	1.2	2.4
	(0.13)	(0.76)	(0.43)
+/- 3 years	-9.6	-2.0	2.6
	(0.00)	(0.71)	(0.38)
+/- 4 years	-14.7	-4.7	1.3
	(0.00)	(0.41)	(0.70)

Table 6: **Average Treatment Effects of SNB's Negative Rates in Isolation**
The table reports the estimated coefficients of the negative rates dummy variable in regression (11) together with their respective p-values (in brackets) obtained using Newey-West standard errors. The first column reports the regression without controls, the second column reports the estimates with the core set of control variables, and the third column contains the results including all the controls. The four samples start in December 2013, December 2012, December 2011, and December 2010, and end in November 2015, November 2016, November 2017, and November 2018, respectively.

2015. Given that such purchases reduce the available stock of government-backed safe assets in the euro area, they could also reduce the relative safety and exclusivity premium of Swiss safe assets.[28] Alternatively, this could have also raised the premia of all safe assets globally, although we stress that any broad-based effects from asset purchases that lower the general interest rate level in the euro area should affect the frictionless yields within our model and not the bond-specific safety premia.

To test this hypothesis, we include ECB government bond holdings divided by the nominal GDP of the euro area in regressions that take a form similar to the two previous exercises:

$$\bar{\Psi}_t = \alpha + \delta_{nr} d_t^{pspp} + \delta_l X_t + \varepsilon_t, \tag{12}$$

where we first use a sample that contains one year of data before the January 22, 2015, launch of the PSPP and one year of data after this date. As before, the regressions are repeated with symmetrically expanded samples that each time contain one more year of data before and after the event date. Given that our sample ends in March 2019, we can add up to four years of data around both sides of the event date, as in the previous section.

The results are reported in Table 7. Overall, they point to a significant *negative* effect on the Swiss safety premium from the ECB's asset purchases, and the effect becomes more significant as we expand the data window around the event date. The magnitude of the effect varies with the length of the sample and the inclusion of controls, but with at least six years of data included the estimate ranges from a -0.4 basis point to a -1.2 basis point per 1 percentage

[28]Koijen et al. (2017) find that foreigners exhibited the strongest reaction to the ECB asset purchases under this program in terms of rebalancing their portfolios toward more attractive investment opportunities elsewhere. That process could also have led them to reassess their perceptions about the relative safety of Swiss safe assets.

Sample	No controls	Core controls	All controls
+/- 1 year	-1.3	-1.4	1.7
	(0.00)	(0.13)	(0.13)
+/- 2 years	-0.8	-1.1	0.1
	(0.00)	(0.02)	(0.91)
+/- 3 years	-1.1	-1.1	-0.4
	(0.00)	(0.00)	(0.30)
+/- 4 years	-1.2	-1.1	-0.6
	(0.00)	(0.00)	(0.00)

Table 7: **Average Treatment Effects of ECB's PSPP in Isolation**
The table reports the estimated coefficients of the PSPP variable in regression (12) together with their respective p-values (in brackets) obtained using Newey-West standard errors. The first column reports the regression without controls, the second column reports the estimates with the core set of control variables, and the third column contains the results including all the controls. The four samples start in January 2014, January 2013, January 2012, and January 2011, and end in December 2015, December 2016, December 2017, and December 2018, respectively.

point of GDP in asset purchases. Given the increase in the ECB's asset holdings through the end of our sample, a point estimate of approximately -1 basis point implies a cumulative effect of close to 20 basis points between spring 2015 and spring 2019. Thus, the decline in the available stock of safe assets in the euro area has materially negatively affected the safety premium commanded by Swiss safe assets. Note also that these results and conclusions are fully consistent with the results from the full sample regressions reported in Table 4.

In terms of tangible evidence from the euro area supportive of our results, we point to Arrata et al. (2019), who find that the asset purchases under the PSPP constrained the free flow of bonds in the highly rated government bond markets in the euro area and caused bonds from those markets to trade at a special premium in repo markets. Furthermore, Roh (2019) documents that this drove up the prices of this class of bonds. These findings suggest the outstanding amount of euro-area safe assets available for trading was significantly reduced, which could reduce the relative exclusiveness of Swiss safe assets and depress their safety premium.

5.4 The Term Structure of Safety Premia

Krishnamurthy and Vissing-Jorgensen (2012) explicitly assume that safety attributes may differ across short- and long-term assets and hence lead to differences in their respective convenience yields. Therefore, after having studied the average safety premium and its determinants, we now turn to the behavior of the term structure of safety premia.

To study this in greater detail, we generate a generic standardized measure of safety premia across maturities by taking the fitted frictionless yield of a given maturity from the AFNS-R model and subtract the fitted yield of the same maturity derived from the flexible

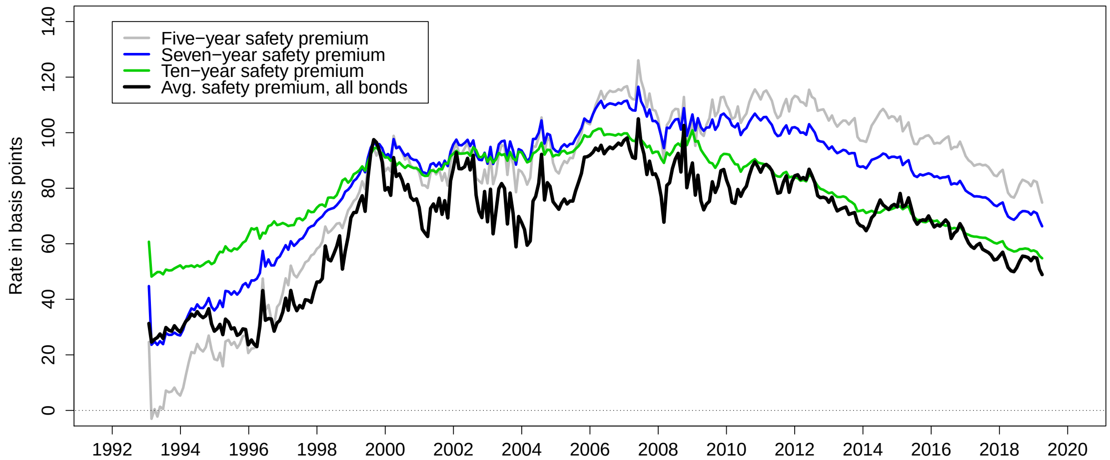


Figure 13: **Term Structure of Swiss Confederation Bond Safety Premia**
Illustration of the estimated term structure of generic safety premia calculated as described in the main text and compared to the average estimated Swiss Confederation bond safety premium implied by the AFNS-R model. The data cover the period from January 29, 1993, to March 31, 2019.

AFGNS model considered in Section 4.4. Since the latter yield just represents a very flexible fit to raw bond price data, this construction essentially provides us with a smoothed measure of the safety premium at a given maturity. This has both advantages and disadvantages. On the positive side, it smooths out some of the variation in the bond-specific premia that tends to be reflected in the average safety premium series and provides us with a consistent measure for fixed maturities. On the negative side, it neglects the compositional changes in the underlying bond universe, such as changes in the number of bonds outstanding, their coupon rates, and remaining time to maturity that may play a role in some of the changes we see in the average safety premium series. Importantly, however, the generic measures are highly positively correlated with the average safety premium analyzed thus far.

In Figure 13, we plot these generic smoothed estimates of the Swiss safety premium at the five-, seven-, and ten-year maturities and compare them to the average estimated Swiss safety premium considered so far. As shown in online Appendix D, the average time to maturity started around nine years in 1993, gradually declined to a low close to five years in 1999, and rapidly increased to around ten years, where it remained until 2012; then started rising to close to fifteen years by the end of our sample. Based on the variation in the average time to maturity for our bond sample, it is therefore not surprising that the average estimated safety premium series started out close to the generic seven-year safety premium, then approached the five-year safety premium by late 1995 and tracked it until 1999. All safety premium series stayed fairly close to each other for the next ten years. However, since 2009 we have started to see a negative term structure develop in which long-term safety premia have declined more than shorter-term safety premia. Furthermore, since the average time to maturity of our bonds is slightly above ten years during most of this period, it is reasonable to see the average safety premium series track the generic ten-year safety premium during this part of the sample.

The negative slope of the term structure of safety premia over the past decade suggests that the very low interest rates in Switzerland since the financial crisis are exerting greater downward pressure on the safety premia of long-term bonds compared to the safety premia of shorter-term bonds. This seems reasonable since the duration risk of very long-term bonds in low and negative interest rate environments is really high. Hence, the variation in the term structure of safety premia would seem to lend support to the view that, provided interest rates become sufficiently low, even the most desirable safe assets can lose some of their attractiveness, in part due to the sizable duration risk their exposure entails.

6 Danish Government Bond Safety Premium Analysis

In this section, we seek to establish further support for our findings for Swiss safety premia by repeating the analysis for a monthly sample of Danish government bond prices covering the period from January 31, 1995, to March 29, 2019. Similar to Swiss Confederation bonds, Danish government bonds are widely viewed as being among the safest government bonds in the world, but lack the liquidity superiority of U.S. Treasuries. Hence, as in our Swiss analysis, we consider any bond-specific premia in their bond prices to reflect safety premia. To estimate these bond-specific premia, we continue to use our parsimonious benchmark AFNS-R model with diagonal $K^{\mathbb{P}}$ and Σ matrices.

Figure 14 shows the average Danish government bond safety premium $\bar{\Psi}_t$ across the outstanding bonds at a given point in time. The average estimated liquidity premium clearly varies notably over time, reaching a minimum of -62 basis points in late 1995 and achieving a maximum of 69 basis points in January 2015 shortly after the SNB had discontinued its minimum exchange rate to the euro and the ECB had announced its first open-ended large-scale purchases of euro area government bonds. For the entire period, the series has an average of 12.14 basis points with a standard deviation of 26.12 basis points.

One notable difference between the Swiss and Danish safety premia is observed during the financial crisis. Events like the onset of the crisis in the summer of 2007 and the bankruptcy of Lehman Brothers in September 2008 coincide with spikes in the Swiss safety premium, while they tend to be associated with a persistent decline in the Danish safety premium. We speculate that Denmark with its peg to the euro was viewed by global investors as less of a safe haven during this period than Switzerland.

Interestingly and more importantly for our analysis, the safety premium appeared to shift persistently upward somewhere in the period between late 1997 and mid-1999. Therefore, the introduction of the euro might have increased the appeal of Danish government bonds as

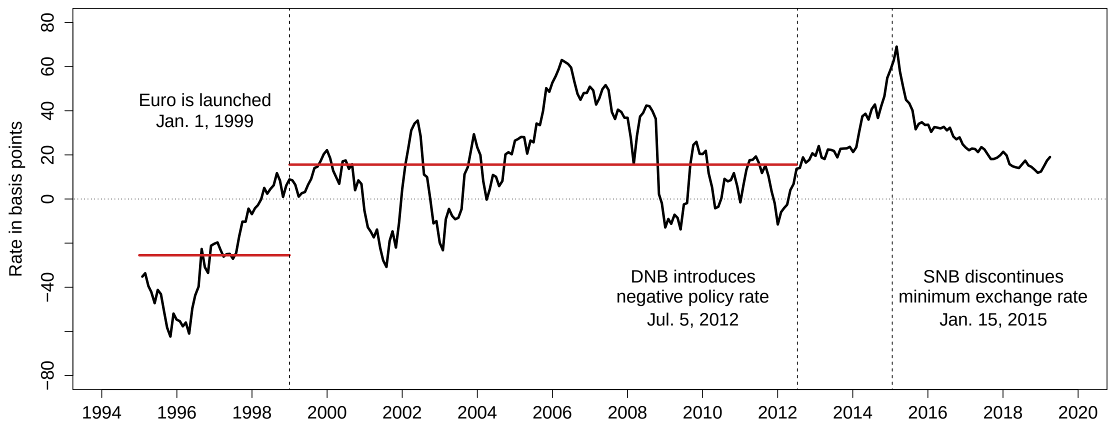


Figure 14: **Average Estimated Danish Government Bond Safety Premium**
Illustration of the average estimated Danish government bond safety premium for each observation date, as implied by the AFNS-R model. The bond safety premiums are measured as the estimated yield difference between the frictionless yield to maturity of individual government bonds with the market risk factor turned off and the corresponding fitted yield to maturity. The data cover the period from January 31, 1995, to January 31, 2019.

safe-haven assets in much the same way as it did for Swiss safe assets. In terms of magnitudes, the Danish safety premium averaged -25.5 basis points from 1995 through December 1998, while its average from January 1999 through June 2012 was 15.6 basis points. Thus, there is a visual upward level shift in the Danish safety premium of more than 40 basis points between these two periods. However, as before, a comprehensive time series analysis is required to establish any connection between the introduction of the euro and the Danish safety premia.

Furthermore, note that the average Danish safety premium appears to have trended higher in the years following the introduction of negative interest rates by the DNB. In an effort to curb the appreciation pressure of the Danish kroner relative to the euro, the DNB lowered its key policy interest rate to -20 basis points in July 2012. In response, the entire yield curve of Danish government bonds shifted downwards as documented in Christensen (2019), and part of this yield decline appears to be driven by upticks in the Danish safety premium. However, to link the introduction of negative interest rates to these trends obviously requires a careful econometric analysis that controls for the effects of confounding factors that may have influenced the safety premium during this period.

Finally, since safe asset purchases by foreign central banks reduce the available stock of foreign safe assets, the government bond purchases made by the ECB since January 2015 through the PSPP could make Danish government bonds relatively less attractive and reduce the Danish safety premium. Alternatively, they could be viewed as reducing the available stock of safe assets globally and hence raise the premium of all safe assets. Which of these two effects is more important is an empirical question we explore as well.

6.1 Regression Analysis

To test the stated hypotheses above, we again first look for structural breaks in the time series of the average estimated Danish safety premium. We run the CUSUM test from Page (1954) to measure the stability of the intercept coefficient in a regression of a vector of ones on the safety premium $\bar{\Psi}_t$. The test shows that there is one significant break in the premium series starting from September 1998 and lasting until the end of the sample. Therefore, the launch of the euro in January 1999 also appears to have given rise to a break in the time series of the Danish safety premium, while this is once more not the case for either the introduction of negative interest rates by the DNB or the ECB's asset purchases.

Similar to the Swiss analysis, we control for a host of confounding factors. In a core set of controls, we consider the VIX, the spread between the Italian and the German 10-year government bond yield, the TED spread, and the on-the-run premium in U.S. Treasuries to proxy for investors' risk aversion, financial market uncertainty, and related demand for safe-haven assets. We consider the spread between German and Danish 10-year government bond yields to proxy for the opportunity cost of holding Danish government bonds and we use the debt-to-GDP ratio to control for effects tied to the supply of Danish government bonds. The overnight deposit rate should proxy for the opportunity cost of holding money and the associated liquidity premia of Danish government bonds. Furthermore, we include the average Danish government bond age and the one-month realized volatility of the ten-year Danish government bond yield as additional proxies for bond liquidity following the work of Houweling et al. (2005). Inspired by the analysis of Hu et al. (2013), we also include a noise measure of Danish government bond prices to control for the variation in the amount of arbitrage capital available in this market.

In addition to the set of core control variables, we again consider several other confounding factors in the regressions. We add the overnight fed funds rate to proxy for the U.S. safe-asset liquidity premium as in Nagel (2016), and use reported earnings per share of companies in the S&P 500 to account for opportunity costs in the equity market. We also consider the MOVE volatility index to proxy for risk aversion in the bond market.

6.1.1 Effects of the Introduction of the Euro in Isolation

To focus narrowly on the impact of the launch of the euro in isolation and avoid potentially polluting effects from the financial crisis, the introduction of negative rates, and the ECB

Sample	No controls	Core controls	All controls
+/- 1 year	6.5	2.5	-0.1
	(0.10)	(0.68)	(0.99)
+/- 2 years	17.6	7.6	2.6
	(0.00)	(0.24)	(0.71)
+/- 3 years	19.8	-1.9	-0.6
	(0.03)	(0.77)	(0.93)
+/- 4 years	29.6	21.4	31.0
	(0.00)	(0.00)	(0.00)

Table 8: **Average Treatment Effects of Euro Introduction in Isolation**
The table reports the estimated coefficients of the euro dummy variable in regression (13) together with their respective *p*-values (in brackets) obtained using Newey-West standard errors. The first column reports the regression without controls, the second column reports the estimates with the core set of control variables, and the third column contains the results including all the controls. The four samples start in January 1998, January 1997, January 1996, January 1995, and end in December 1999, December 2000, December 2001, and December 2002, respectively.

asset purchases, we repeat our simplified regressions of the form:[29]

$$\bar{\Psi}_t = \alpha + \delta_{euro} d_t^{euro} + \delta_l X_t + \varepsilon_t, \tag{13}$$

where we first use a sample that contains one year of data before the January 1, 1999, euro launch and one year of data after this date. We then expand this sample in a symmetric fashion by adding one more year of data before and after the euro launch and rerun the regression. This exercise is repeated until the sample contains four years of data before and after the euro launch, which is the maximum given the January 1995 start date for our Danish bond price data. Overall, this allows us to compare the residual variation in the immediate pre-euro period to that during the immediate post-launch period.

The results are reported in Table 8. For the shorter samples with up to three years of data before and after the euro launch, the results are erratic and insignificant, which may be due to low statistical power. However, for the most informative sample with four years of data before and after the euro launch, the estimated dummy coefficients are relatively stable, systematically highly statistically significant, and not very sensitive to the set of controls used. Therefore, we see clear evidence of a statistically significant *positive* effect on the Danish safety premium following the introduction of the euro of roughly 20-30 basis points.

One might view this result as surprising given that investors knew the launch date well in advance and therefore could have rebalanced their portfolios towards Danish assets earlier. Our results suggest that the effect starts to be reflected in the safety premium a couple of years before the launch and is fully priced in a couple of years after it. As in the Swiss

[29] In these regressions, we again have to drop the lags of the explanatory variables due to the limited number of observations and high number of control variables.

Sample	No controls	Core controls	All controls
+/- 1 year	16.3	3.5	8.6
	(0.00)	(0.54)	(0.29)
+/- 2 years	16.8	10.1	16.5
	(0.00)	(0.03)	(0.00)
+/- 3 years	24.4	2.8	2.0
	(0.00)	(0.54)	(0.71)
+/- 4 years	25.5	6.1	10.0
	(0.00)	(0.17)	(0.05)

Table 9: **Average Treatment Effects of the DNB's Negative Rates in Isolation**
The table reports the estimated coefficients of the negative rates dummy variable in regression (14) together with their respective p-values (in brackets) obtained using Newey-West standard errors. The first column reports the regression without controls, the second column reports the estimates with the core set of control variables, and the third column contains the results including all the controls. The four samples start in July 2011, July 2010, July 2009, and July 2008, and end in June 2013, June 2014, June 2015, and June 2016, respectively.

case, this seems reasonable given that the involved portfolio flows likely were informed by the general experience with the euro and the new common monetary policy in the euro area.

6.1.2 Effects of the Introduction of Negative Rates in Isolation

To assess the impact of the effects from introducing negative rates in isolation, we follow a similar approach and run simplified regressions of the following form:

$$\bar{\Psi}_t = \alpha + \delta_{nr} d_t^{nr} + \delta_l X_t + \varepsilon_t, \tag{14}$$

where we first use a sample that contains one year of data before the July 5, 2012, introduction of negative rates and one year of data after this date. As before, the regressions are repeated with symmetrically expanded samples that each time contain one more year of data before and after the event date. For consistency with the previous exercise, we add up to four years of data around both sides of the event date.

The results are reported in Table 9, in which we note a significant and large positive effect on the Danish safety premium from the introduction of negative interest rates when we exclude controls. Once they are included, the magnitude of the estimated coefficients is significantly diminished, and their significance mostly vanishes. Similar to what we found for Switzerland, any effects on the Danish safety premium during this period therefore appear to be due to other factors.

Sample	No controls	Core controls	All controls
+/- 1 year	-3.4	-11.2	-16.6
	(0.01)	(0.02)	(0.02)
+/- 2 years	-0.9	-8.0	-7.0
	(0.10)	(0.00)	(0.00)
+/- 3 years	-0.6	-5.8	-6.1
	(0.14)	(0.00)	(0.00)
+/- 4 years	-0.4	-3.5	-4.6
	(0.21)	(0.00)	(0.00)

Table 10: **Average Treatment Effects of ECB's PSPP in Isolation**
The table reports the estimated coefficients of the PSPP variable in regression (15) together with their respective p-values (in brackets) obtained using Newey-West standard errors. The first column reports the regression without controls, the second column reports the estimates with the core set of control variables, and the third column contains the results including all the controls. The four samples start in January 2014, January 2013, January 2012, and January 2011, and end in December 2015, December 2016, December 2017, and December 2018, respectively.

6.1.3 Effects of the ECB's PSPP in Isolation

One particular factor we are interested in exploring further is the effect of the ECB's government bond purchases on the Danish safety premium. Given that such purchases reduce the available stock of government-backed safe assets in the euro area, they could potentially reduce the relative safety and exclusivity premium of Danish safe assets similar to what we observe for the Swiss safety premia.

To test the stated hypotheses, we include the ECB government bond holdings divided by the nominal GDP of the euro area in regressions that take a form similar to the two previous exercises:

$$\bar{\Psi}_t = \alpha + \delta_{nr} d_t^{pspp} + \delta_l X_t + \varepsilon_t, \quad (15)$$

where we first use a sample that contains one year of data before the January 22, 2015, launch of the PSPP and one year of data after this date. As before, the regressions are repeated with symmetrically expanded samples that each time contain one more year of data before and after the event date. Given that our sample ends in March 2019, we can add up to four years of data around both sides of the event date as in the exercises in the previous sections.

The results are reported in Table 10. Overall, they point to a significant and sizable *negative* effect on the Danish safety premium from the ECB's asset purchases. The effect becomes more significant as we expand the data window around the January 22, 2015, event date. The magnitude of the effect varies with the length of the sample and the inclusion of controls, but with at least six years of data and controls included the estimate ranges from -3.5 basis points to -6.1 basis points per percentage point of GDP in asset purchases. Given the increase in the ECB's asset holdings through the end of our sample, a point estimate of

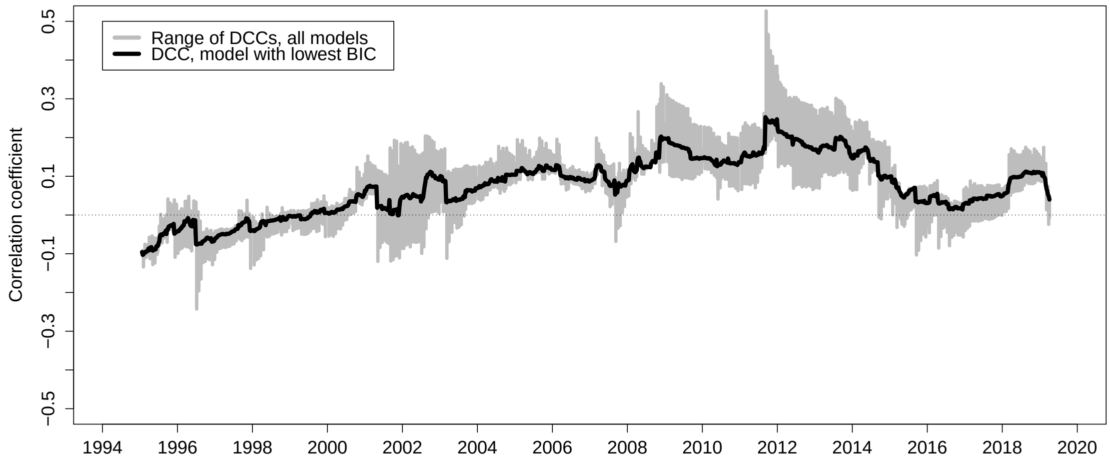


Figure 15: **Dynamic Conditional Correlations of Swiss and Danish Safety Premia**
Illustration of the average estimated Danish government bond safety premium for each observation date, as implied by the AFNS-R model. The bond safety premiums are measured as the estimated yield difference between the frictionless yield to maturity of individual government bonds with the market risk factor turned off and the corresponding fitted yield to maturity. The data cover the period from January 31, 1995, to January 31, 2019.

around -4 basis points implies a cumulative effect of close to 80 basis points between spring 2015 and spring 2019. Thus, the declines in the available stock of safe assets in the euro area has materially negatively affected the safety premium commanded by Danish safe assets.

The estimated effects are qualitatively similar to those found for Swiss safety premia. However, the magnitude is an order of magnitude larger. The difference in sensitivity is likely tied to the fact that Denmark is a member of the European Union and that the Danish kroner is pegged to the euro. Thus, the Danish financial markets are extremely closely tied to those in the euro area. In contrast, Switzerland is not a member of the European Union, and the Swiss franc has been freely floating against the euro since January 2015, creating some room for the exchange rate to adjust in response to adjustments in Swiss bond premia.

6.2 Dynamic Correlations between Swiss and Danish Safety Premia

As we have shown in the previous sections, both Swiss and Danish safety premia reacted significantly to the introduction of the euro and to the ECB's asset purchases. In this section, we therefore examine their co-movement and analyze whether these events might have affected their dynamic relationship.

To begin, we look at the sample correlation of first differences in the safety premia in three subsamples: the period before the introduction of the euro (January 1995-December 1998), the period after the euro introduction but before the start of the ECB's asset purchase program (January 1999-December 2014), and the period after the ECB's asset purchases started (January 2015-March 2019). We find that the correlations in the three periods are -0.05, 0.15, and -0.08, respectively. The difference between the correlations in the first and the second period, and between those in the second and third period are statistically significant, with Fisher's z-score of -2.49 and 2.95, respectively.

We also estimate a time-varying correlation coefficient between the first differences in the Swiss and Danish safety premia by using a version of the general autoregressive conditional heteroskedasticity (GARCH) model that allows for dynamic conditional correlation (DCC) as proposed by Engle (2002). Figure 15 plots the time-varying correlations estimated using a variety of DCC GARCH (m,n) models, in which m = 1,2,3 and n = 1,2,3 (gray-shaded area) and the DCC GARCH (1,1) model with the lowest BIC (solid black line).

The correlation between the safety premia turned positive around the time of the introduction of the euro and reached its maximum around the time of the peak of the European sovereign debt crisis in the summer of 2012. After the ECB started to purchase government bonds in the euro area in January 2015, the correlation trended lower but recovered somewhat toward the end of the sample. This is a pattern that would be consistent with a gradual decline in the importance of safety premia in Swiss and Danish government bond yields since 2015, as also indicated by our regression analysis.

To summarize, the analysis of the Danish safety premium produces results that are very similar to our findings for the Swiss safety premium. First and most importantly, the launch of the euro appears to have boosted Danish safety premia by 20-30 basis points. Second, the Danish safety premia were not affected by the introduction of negative rates. Last, the ECB's purchases of safe government bonds from the euro area have put significant downward pressure on Danish safety premia. Therefore, we take the presented evidence to provide full support for the conclusions we draw from the analysis of the Swiss safety premium.

7 Conclusion

In this paper, we are among the first to estimate bond-specific premia, frequently labeled convenience yields, in a conventional government bond market, specifically the market for Swiss Confederation bonds. Since these bonds are much less liquid than U.S. Treasuries, we attribute their premia to the high safety and creditworthiness of the Swiss Confederation and therefore refer to them as *safety premia*. To support this interpretation in our empirical analysis, we control for a large number of factors, including investor risk aversion as well as liquidity and debt supply effects.

We find a sizable safety premium of Swiss safe assets averaging 68 basis points and exhibiting significant variation over time. One important implication of this result is that the yields of Swiss Confederation bonds are below those we would observe in a world without excess demand tied to safety concerns—a point also made by Krishnamurthy and Vissing-Jorgensen

(2012) in the context of the U.S. Treasury market.

In addition to describing the variation of the Swiss safety premium, we also explore its relationship to three key events over the past 25 years: the launch of the euro in January 1999, the introduction of negative interest rates in Switzerland in December 2014, and the ECB's launch of its main asset purchase program, the PSPP, in January 2015.

Based on a range of tests of a break in the Swiss safety premium in January 1999, we conclude that the launch of the euro led to a long-lasting upward shift in the Swiss safety premium of between 35 and 40 basis points. We conjecture that this effect came about because, once the euro was created, even German and French bonds could be exposed to some extreme tail risk such as a bailout of one or more euro member states, which would leave them seriously indebted. All else being equal, such scenarios would increase the attractiveness of safe assets in a safe-haven country such as Switzerland.

More recently, a combination of negative interest rates and ECB asset purchases appear to have given rise to a persistent negative pressure on the Swiss safety premium. Negative effects from these two developments could be expected. A move to negative rates should make government debt less attractive in general. Furthermore, safe asset purchases by foreign central banks reduce the supply of foreign safe assets and, in turn, reduce the relative uniqueness of Swiss Confederation bonds within the class of highly safe European assets. These conclusions are supported by matching evidence from Danish government bond prices, which represent another class of extremely safe assets. Our results hence point to a potentially important international spillover channel of central bank asset purchases that operates through its impact on the relative scarcity of safe assets. These findings also raise the prospect that even the safest assets can become unattractive when interest rates are sufficiently low. To shed more light on that question, it may be relevant to look at other countries with very safe government debt and negative interest rates. However, we leave it for future research to explore those avenues.

Appendix

In this appendix, we first specify the risk-neutral $\mathbb{Q}$-dynamics and the objective $\mathbb{P}$-dynamics in the AFNS-R-L model introduced in Section 5.2.2 before we proceed to a description of the construction of our observable composite liquidity factor.

In general, the short rate used for discounting bond cash flows in the AFNS-R-L model takes the form

$$\overline{r}_t^i = L_t + S_t + \beta^i(1 - e^{-\lambda^{R,i}(t-t_0^i)})X_t^R + \delta^L X_t^{liq},$$

where the risk-neutral $\mathbb{Q}$-dynamics of the state variables used for pricing are given by

$$\begin{pmatrix} dL_t \\ dS_t \\ dC_t \\ dX^R \\ dX_t^{liq} \end{pmatrix} = \begin{pmatrix} 0 & 0 & 0 & 0 & 0 \\ 0 & \lambda & -\lambda & 0 & 0 \\ 0 & 0 & \lambda & 0 & 0 \\ 0 & 0 & 0 & \kappa_R^{\mathbb{Q}} & 0 \\ 0 & 0 & 0 & 0 & \kappa_{liq}^{\mathbb{Q}} \end{pmatrix} \left[\begin{pmatrix} 0 \\ 0 \\ 0 \\ \theta_R^{\mathbb{Q}} \\ \theta_{liq}^{\mathbb{Q}} \end{pmatrix} - \begin{pmatrix} L_t \\ S_t \\ C_t \\ X_t^R \\ X_t^{liq} \end{pmatrix} \right] dt + \Sigma \begin{pmatrix} dW_t^{L,\mathbb{Q}} \\ dW_t^{S,\mathbb{Q}} \\ dW_t^{C,\mathbb{Q}} \\ dW_t^{R,\mathbb{Q}} \\ dW_t^{liq,\mathbb{Q}} \end{pmatrix}$$

and Σ continues to be a diagonal matrix.

To complete the description of the AFNS-R-L model, we specify an essentially affine risk premium structure, which implies that the risk premia Γ_t take the form

$$\Gamma_t = \gamma^0 + \gamma^1 X_t,$$

where $\gamma^0 \in \mathbf{R}^5$ and $\gamma^1 \in \mathbf{R}^{5\times 5}$ contain unrestricted parameters. Thus, the resulting unrestricted AFNS-R-L model has $\mathbb{P}$-dynamics given by

$$\begin{pmatrix} dL_t \\ dS_t \\ dC_t \\ dX_t^R \\ dX_t^{liq} \end{pmatrix} = \begin{pmatrix} \kappa_{11}^{\mathbb{P}} & \kappa_{12}^{\mathbb{P}} & \kappa_{13}^{\mathbb{P}} & \kappa_{14}^{\mathbb{P}} & \kappa_{15}^{\mathbb{P}} \\ \kappa_{21}^{\mathbb{P}} & \kappa_{22}^{\mathbb{P}} & \kappa_{23}^{\mathbb{P}} & \kappa_{24}^{\mathbb{P}} & \kappa_{25}^{\mathbb{P}} \\ \kappa_{31}^{\mathbb{P}} & \kappa_{32}^{\mathbb{P}} & \kappa_{33}^{\mathbb{P}} & \kappa_{34}^{\mathbb{P}} & \kappa_{35}^{\mathbb{P}} \\ \kappa_{41}^{\mathbb{P}} & \kappa_{42}^{\mathbb{P}} & \kappa_{43}^{\mathbb{P}} & \kappa_{44}^{\mathbb{P}} & \kappa_{45}^{\mathbb{P}} \\ \kappa_{51}^{\mathbb{P}} & \kappa_{52}^{\mathbb{P}} & \kappa_{53}^{\mathbb{P}} & \kappa_{54}^{\mathbb{P}} & \kappa_{55}^{\mathbb{P}} \end{pmatrix} \left(\begin{pmatrix} \theta_1^{\mathbb{P}} \\ \theta_2^{\mathbb{P}} \\ \theta_3^{\mathbb{P}} \\ \theta_4^{\mathbb{P}} \\ \theta_5^{\mathbb{P}} \end{pmatrix} - \begin{pmatrix} L_t \\ S_t \\ C_t \\ X_t^R \\ X_t^{liq} \end{pmatrix} \right) dt + \Sigma \begin{pmatrix} dW_t^{L,\mathbb{P}} \\ dW_t^{S,\mathbb{P}} \\ dW_t^{C,\mathbb{P}} \\ dW_t^{R,\mathbb{P}} \\ dW_t^{liq,\mathbb{P}} \end{pmatrix}.$$

This is the transition equation in the Kalman filter estimation.

Construction of the Observable Composite Liquidity Factor

In this section, we describe the construction of the observable composite liquidity factor. It has three components, each of which is assumed to contain important, but incomplete information about the liquidity risk of the Swiss Confederation bond market. Thus, by combining the three proxies for bond liquidity risk, we hope to iron out idiosyncratic noise in addition to avoid being overly dependent on any single measure.

First, inspired by the analysis of Hu et al. (2013), we include a noise measure of Swiss

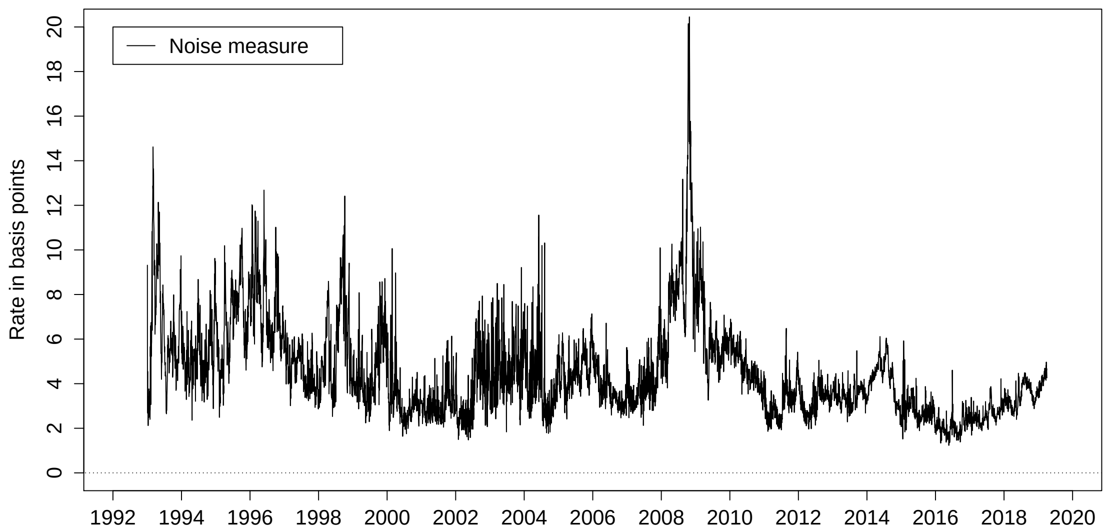


Figure 16: **Noise Measure**
Illustration of the Swiss noise measures constructed based on the fitted errors implied by the estimated AFGNS model using the Swiss Confederation bonds covering the period from January 4, 1993 to March 29, 2019.

Confederation bond prices to capture variation in the amount of arbitrage capital available in the market for these bonds. In principle, this could be constructed using any yield curve model. However, we choose to focus on the very flexible five-factor arbitrage-free generalized Nelson-Siegel (AFGNS) model developed in Christensen et al. (2009), which we estimate with the one-step approach recommended by Andreasen et al. (2019).[30] The resulting noise measure defined as the average absolute fitted errors of the yield to maturity across all available bonds at each observation date is shown in Figure 16.

Second, we note that Houweling et al. (2005) recommend to use yield volatility as a proxy for bond liquidity. To that end, we use a standard measure of realized yield volatility based on daily data. First, we again rely on the output from the AFGNS model estimated using the daily sample of Swiss Confederation bond prices described above. This gives us daily fitted zero-coupon yields at all relevant maturities. We then generate the realized standard deviation of daily changes in interest rates for the past 31-day period on a rolling basis. The realized variance measure is used by Andersen and Benzoni (2010), Collin-Dufresne et al. (2009), as well as Jacobs and Karoui (2009) in their assessments of stochastic volatility models. This measure is fully nonparametric and has been shown to converge to the underlying realization of the conditional variance as the sampling frequency increases; see Andersen et al. (2003) for details. The square root of this measure retains these properties. For each observation date

[30]We stress that the results are robust to using the Svensson (1995) model, see online Appendix D.

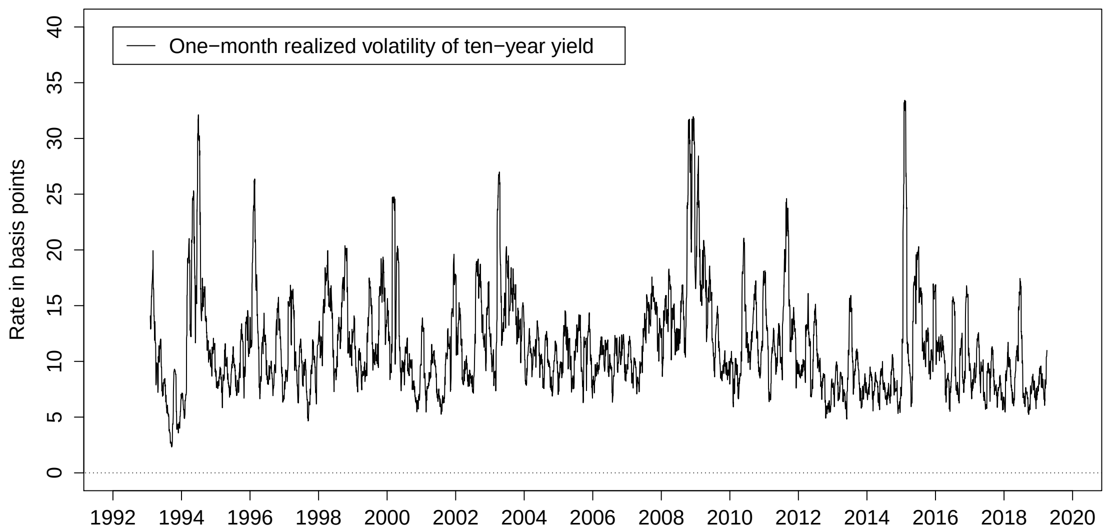


Figure 17: **Realized Yield Volatility Measure**
The top panel shows the one-month realized volatility of yields with three different maturities, all constructed from the AFGNS model estimated with our daily sample of Swiss Confederation bond prices. The bottom panel shows the one-month realized volatility of Swiss ten-year yields constructed from three different yield curve models, all estimated with our daily sample of Swiss Confederation bond prices.

t we determine the number of trading days N during the past 31-day time window (where N is most often 21 or 22).[31] We then generate the realized standard deviation as

$$RV_{t,\tau}^{STD} = \sqrt{\sum_{n=1}^{N} \Delta y_{t+n/N}^2(\tau)},$$

where $\Delta y_{t+n/N}(\tau)$ is the change in yield $y(\tau)$ from trading day $(n-1)$ to trading day n.[32]

Figure 17 shows the realized ten-year yield volatility series used in our liquidity factor construction. We focus on the ten-year maturity since the average time to maturity of the available Swiss Confederation bonds is close to ten years for much of our sample period. However, we stress that our results are not sensitive to this particular choice or to the yield curve model used to construct the fitted zero-coupon yields, as shown in online Appendix D.

As the final proxy for Swiss government bond liquidity, we use bid and ask prices for all bonds in our sample kindly provided by staff at the Swiss National Bank. We note that this information has only been collected since January 1998, which explains the later

[31] As a consequence, the realized volatility measure can be calculated for the period from February 4, 1993, to March 29, 2019.

[32] Note that other measures of realized volatility have been used in the literature, such as the realized mean absolute deviation measure as well as fitted GARCH estimates.

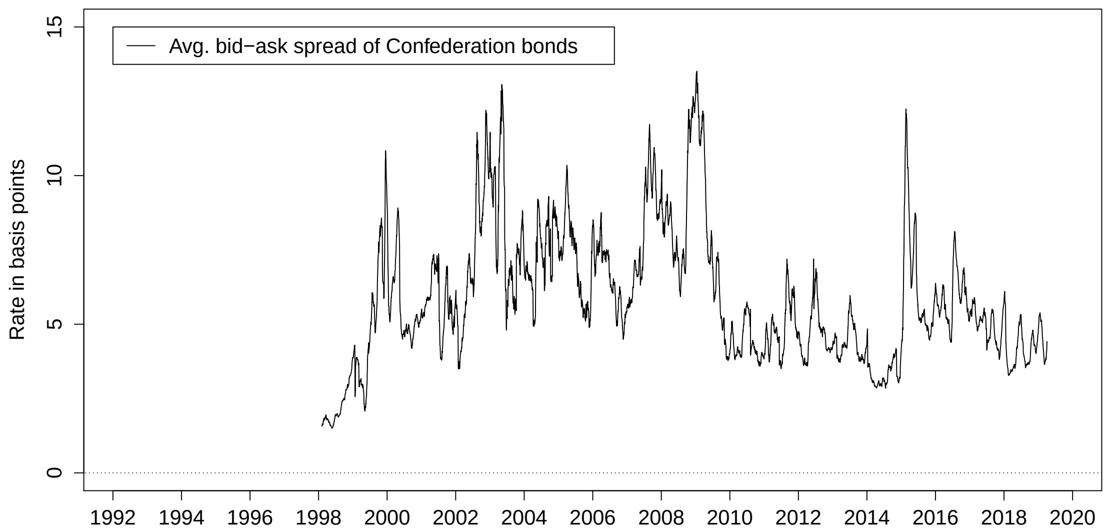


Figure 18: **Average Bid-Ask Yield Spread of Swiss Confederation Bonds**
Illustration of the average bid-ask yield spread of the Swiss Confederation bonds in our sample measured in basis points and constructed as described in the text. The shown series covers the period from February 5, 1998, to March 29, 2019.

stating point for this series. We convert the bid and ask prices into their corresponding yield to maturity equivalents. This allows us to calculate bid-ask yield spreads measured in basis points for each bond in our sample. We then clean the data in three steps. First, we eliminate all observations with either a negative bid-ask spread or a bid-ask spread above 50 basis points—we consider these observations to reflect reporting errors rather than true observations. Second, we eliminate the bid-ask spreads of bonds with less than three months to maturity similar to what we do in the model estimations. Finally, we smooth each individual bond bid-ask yield spread series by calculating the four-week moving average.[33] Once we have this cleaned panel of individual bid-ask yield spread series, we focus on the average bid-ask spread of the available bonds for each observation date, which is shown in Figure 18.

To construct our composite liquidity factor, we follow the approach of Abrahams et al. (2016) and standardize all three variables before calculating the average. To smooth the series further and make it closer to the yield data used in the model estimation, we calculate the four-week moving average. To align the data with our monthly yield data, we focus on the value at the end of each month and deduct the minimum value of this series to eliminate negative values. We define the resulting series shown in Figure 10 in the paper as our monthly observable composite liquidity factor.

[33]As a consequence, our bid-ask yield spread measure can be calculated for the period from February 5, 1998, to March 29, 2019.

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Online Appendix

The Safety Premium of Safe Assets

Jens H. E. Christensen
Federal Reserve Bank of San Francisco
jens.christensen@sf.frb.org

&

Nikola Mirkov
Visiting researcher, ITAM
nikola.mirkov@gmail.com

The views in this paper are solely the responsibility of the authors and should not be interpreted as reflecting the views of the Federal Reserve Bank of San Francisco or the Federal Reserve System.

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Contents

A Swiss Confederation Bond Data

This appendix provides the details of the Swiss Confederation bonds considered in the analysis.

Confederation bond	No. obs.	Issuance		Number of auctions	Total amount
		Date	Amount		
(1) 6.75% 1/22/1999	69	1/22/1991	324	2	824
(2) 6.25% 3/15/1999	71	3/15/1991	206	1	206
(3) 6.25% 7/15/1999	74	5/17/1991	497	1	497
(4) 6.25% 7/15/2000	87	7/15/1991	302	1	302
(5) 6.5% 2/5/2002	106	2/5/1992	600	2	1,400
(6) 6.5% 4/10/2004	132	4/10/1992	704	1	704
(7) 6.75% 6/11/2003	122	6/11/1992	1,205	1	1,205
(8) 7% 7/9/2001	99	7/9/1992	500	2	1,000
(9) 7% 9/10/2003	124	9/10/1992	700	1	700
(10) 6.25% 11/5/2004	139	11/5/1992	700	1	700
(11) 6.25% 1/7/2003	117	1/7/1993	998	2	1,609
(12) 5.25% 2/11/1998	57	2/11/1993	800	2	1,154
(13) 5% 3/11/2000	81	3/11/1993	500	5	1,948
(14) 4.5% 4/8/2006	153	4/8/1993	477	10	4,291
(15) 4.5% 6/10/2000	81	6/10/1993	550	5	2,351
(16) 4.5% 7/8/2002	105	7/8/1993	231	9	2,858
(17) 4.5% 10/7/2004	129	10/7/1993	862	6	3,816
(18) 4.25% 1/6/2014	237	1/6/1994	676	8	3,208
(19) 4% 3/10/1999	56	3/10/1994	353	3	907
(20) 5.5% 10/7/2001	81	10/7/1994	685	2	1,140
(21) 5% 11/10/1996	21	11/10/1994	555	1	555
(22) 5.5% 1/6/2005	117	1/6/1995	726	3	1,769
(23) 4.25% 1/8/2008	141	1/8/1996	1,000	7	5,366
(24) 4.5% 6/10/2007	129	6/10/1996	302	7	4,468
(25) 4.25% 6/5/2017	237	6/5/1997	228	10	3,260
(26) 3.5% 8/7/2010	153	8/7/1997	561	10	8,902
(27) 3.25% 2/11/2009	118	2/11/1998	938	8	8,523
(28) 4% 2/11/2023	243	2/11/1998	231	9	2,858
(29) 4% 4/8/2028	243	4/8/1998	467	10	3,012
(30) 2.75% 6/10/2012	149	6/10/1999	569	13	6,360
(31) 4% 1/6/2049	234	1/6/1999	189	9	1,595
(32) 4% 2/11/2013	151	2/11/2000	586	9	5,380
(33) 4% 6/10/2011	128	6/13/2000	1,332	7	6,182
(34) 3.75% 6/10/2015	164	6/11/2001	524	7	3,139
(35) 3% 1/8/2018	174	1/8/2003	1,139	11	6,136
(36) 2.5% 3/12/2016	152	3/12/2003	715	11	5,704
(37) 3.5% 4/8/2033	192	4/8/2003	486	7	2,433
(38) 3% 5/12/2019	177	5/12/2004	309	10	5,399
(39) 1.75% 11/5/2009	57	11/5/2004	628	3	1,756
(40) 2.25% 7/6/2020	165	7/6/2005	423	11	4,101
(41) 2% 10/12/2016	128	10/12/2005	483	6	2,667
(42) 2% 11/9/2014	105	11/9/2005	547	3	1,606
(43) 2.5% 3/8/2036	157	3/8/2006	222	9	2,303
(44) 3.25% 6/27/2027	142	6/27/2007	101	8	1,909
(45) 2% 4/28/2021	108	4/28/2010	675	11	3,958
(46) 2% 5/25/2022	95	5/25/2011	531	8	3,233
(47) 2.25% 6/22/2031	94	6/22/2011	478	7	1,920
(48) 1.5% 4/30/2042	83	4/30/2012	872	8	3,196
(49) 1.25% 6/11/2024	81	6/11/2012	216	11	2,743
(50) 1.25% 6/27/2037	81	6/27/2012	267	11	3,049
(51) 1.5% 7/24/2025	68	7/24/2013	540	8	2,467
(52) 1.25% 5/28/2026	58	5/28/2014	301	13	2,062
(53) 2% 6/25/2064	57	6/25/2014	565	5	1,487
(54) 0.5% 5/27/2030	46	5/27/2015	318	10	1,798
(55) 0.5% 5/30/2058	34	5/30/2016	169	8	1,208
(56) 0% 6/22/2029	33	6/22/2016	308	11	2,213
(57) 0.5% 5/24/2055	22	5/24/2017	80	7	1,102
(58) 0.5% 6/28/2045	21	6/28/2017	255	7	999
(59) 0.5% 6/27/2032	9	6/27/2018	206	5	664

Table 1: **The Universe of Swiss Confederation Bonds**
The table reports the characteristics, issuance dates, initial issuance amounts, total number of auctions, and total issuance amounts in millions of Swiss francs for each Confederation bond. Also reported are the number of monthly observation dates for each bond during the sample period from January 29, 1993, to March 29, 2019.

B Model Fit and Bond-Specific Risk Parameter Estimates

The fit of the AFNS and AFNS-R models along with the estimated bond-specific risk sensitivity parameters in the latter model for each Swiss Confederation bond are given in Table 2.

The first two columns in Table 2 show that the bond pricing errors produced by the AFNS model indicate a reasonable fit, with an overall RMSE of 9.08 basis points. The following two columns reveal a substantial improvement in the pricing errors when correcting for the bond-specific risk premia, because the AFNS-R model has a much lower overall RMSE of just 5.78 basis points. Hence, accounting for these risk premia leads to a notable improvement in the ability of our model to explain the confederation bond prices. Without exception there is uniform improvement in model fit from augmenting the AFNS model.

The final four columns of Table 2 report the estimates of the bond-specific parameters associated with each Confederation bond. Except for a handful of bonds, including one fifty-year bond (number 53), most bonds in our sample are exposed to bond-specific risk as their β^i parameters are large and statistically significant.

Confederation bond	Pricing errors				Estimated parameters			
	AFNS		AFNS-R		AFNS-R			
	Mean	RMSE	Mean	RMSE	β^i	SE	$\lambda^{R,i}$	SE
(1) 6.75% 1/22/1999	10.76	16.84	-1.33	11.18	2.3133	0.1703	9.9982	1.6093
(2) 6.25% 3/15/1999	5.32	17.22	-0.06	6.33	3.9044	0.2841	9.9934	1.0741
(3) 6.25% 7/15/1999	-6.55	18.40	0.49	10.21	2.4130	0.1626	1.2653	0.2087
(4) 6.25% 7/15/2000	2.41	17.83	2.81	14.76	1.2629	0.0903	1.0618	0.2000
(5) 6.5% 2/5/2002	4.08	8.92	0.40	4.97	1.3750	0.3164	0.2498	0.2159
(6) 6.5% 4/10/2004	1.73	8.33	0.80	5.69	2.3906	0.5216	0.0946	0.0354
(7) 6.75% 6/11/2003	2.39	9.24	0.56	6.02	39.2904	0.8659	0.0039	0.0003
(8) 7% 7/9/2001	1.79	7.72	0.47	5.62	1.2626	0.0916	0.6938	0.1995
(9) 7% 9/10/2003	3.23	12.35	-0.09	10.29	72.5920	0.8295	0.0020	0.0001
(10) 6.25% 11/5/2004	4.80	9.66	0.43	6.74	1.5001	0.1448	0.2605	0.0863
(11) 6.25% 1/7/2003	1.90	8.73	0.67	5.67	23.4727	0.8298	0.0070	0.0005
(12) 5.25% 2/11/1998	-3.66	13.70	0.43	8.14	24.0735	0.8213	0.0057	0.0005
(13) 5% 3/11/2000	-4.93	11.38	-0.24	8.09	38.5223	0.8627	0.0044	0.0004
(14) 4.5% 4/8/2006	-1.52	6.64	-0.29	4.50	1.6162	0.1273	0.2928	0.0922
(15) 4.5% 6/10/2000	-5.83	10.76	-1.14	7.41	24.6930	0.8188	0.0074	0.0006
(16) 4.5% 7/8/2002	-3.16	9.28	-0.86	5.33	3.5347	0.7954	0.0638	0.0200
(17) 4.5% 10/7/2004	-1.50	5.90	-0.27	4.48	1.6574	0.1694	0.2737	0.0967
(18) 4.25% 1/6/2014	1.93	7.60	0.27	4.43	76.7614	0.8264	0.0014	0.0001
(19) 4% 3/10/1999	-9.20	13.70	-1.50	8.43	140.4379	0.8460	0.0018	0.0002
(20) 5.5% 10/7/2001	0.47	10.36	0.02	8.08	1.5722	0.5510	0.2594	0.2234
(21) 5% 11/10/1996	-3.12	16.11	0.60	12.08	22.438	0.8986	0.0078	0.0008
(22) 5.5% 1/6/2005	0.01	6.18	0.27	4.89	1.6485	0.1538	0.3039	0.0828
(23) 4.25% 1/8/2008	-1.89	7.62	0.24	4.04	1.6608	0.1141	0.4256	0.1389
(24) 4.5% 6/10/2007	-2.10	7.23	0.05	4.17	1.5642	0.1005	2.1419	0.8660
(25) 4.25% 6/5/2017	0.07	5.88	0.03	4.01	7.3577	0.8925	0.0178	0.0027
(26) 3.5% 8/7/2010	0.17	9.76	0.33	6.34	1.8668	0.1264	0.2807	0.0552
(27) 3.25% 2/11/2009	3.58	13.36	1.11	6.45	1.6901	0.1119	0.5472	0.2007
(28) 4% 2/11/2023	3.43	7.96	0.09	4.54	1.7893	0.2657	0.0832	0.0271
(29) 4% 4/8/2028	4.62	8.86	0.51	6.01	1	n.a.	8.4977	0.9888
(30) 2.75% 6/10/2012	-2.70	8.01	-0.06	5.07	2.3356	0.2128	0.1687	0.0399
(31) 4% 1/6/2049	3.36	11.24	0.00	6.32	0.3337	0.0653	9.9873	1.0561
(32) 4% 2/11/2013	0.83	6.66	-0.02	4.03	2.2104	0.2141	0.1863	0.0490
(33) 4% 6/10/2011	1.67	7.32	-0.02	4.92	1.9441	0.1462	0.3257	0.0890
(34) 3.75% 6/10/2015	-1.32	5.07	0.21	4.01	2.6417	0.3108	0.1257	0.0305
(35) 3% 1/8/2018	-0.88	5.56	0.04	3.96	2.4043	0.2640	0.1401	0.0337
(36) 2.5% 3/12/2016	-3.38	5.72	0.15	4.21	2.3993	0.2047	0.1891	0.0405
(37) 3.5% 4/8/2033	-1.36	7.25	-0.82	6.03	0.8221	0.0270	9.9921	1.4182
(38) 3% 5/12/2019	-1.61	5.95	0.00	4.19	2.2270	0.1962	0.1668	0.0328
(39) 1.75% 11/5/2009	-12.72	27.22	-0.28	14.05	164.9121	1.7678	0.0038	0.0002
(40) 2.25% 7/6/2020	-1.02	4.86	0.18	4.00	1.9344	0.1404	0.2366	0.0545
(41) 2% 10/12/2016	-4.52	6.99	0.19	4.44	2.2048	0.1348	0.3705	0.0890
(42) 2% 11/9/2014	-8.22	13.51	0.19	7.68	2.3374	0.1194	0.4899	0.0764
(43) 2.5% 3/8/2036	-4.49	8.08	-0.65	5.04	0.7164	0.0372	9.9993	1.4149
(44) 3.25% 6/27/2027	3.34	6.67	-0.06	4.49	1.1105	0.0250	9.9867	1.6540
(45) 2% 4/28/2021	-0.01	3.97	0.09	2.96	1.7559	0.1293	0.6022	0.3185
(46) 2% 5/25/2022	1.07	3.71	0.09	2.69	1.6147	0.1378	0.7530	0.6484
(47) 2.25% 6/22/2031	0.42	3.34	0.01	2.71	0.8866	0.0646	9.9845	6.3796
(48) 1.5% 4/30/2042	-7.28	8.89	-0.13	2.86	0.5482	0.0740	9.9939	7.5394
(49) 1.25% 6/11/2024	0.66	3.53	0.07	2.07	1.4239	0.1470	0.7144	0.8508
(50) 1.25% 6/27/2037	-5.27	6.55	-0.06	2.75	0.6711	0.0787	9.9980	7.7019
(51) 1.5% 7/24/2025	1.32	3.69	0.04	2.01	1.3203	0.1782	0.6871	1.1949
(52) 1.25% 5/28/2026	-0.25	2.38	0.05	2.25	1.2600	0.1840	0.7983	1.5809
(53) 2% 6/25/2064	3.52	6.35	0.07	4.28	0.0000	0.1051	9.9897	13.1894
(54) 0.5% 5/27/2030	-2.92	3.43	-0.04	1.98	0.9754	0.1195	8.9381	14.7901
(55) 0.5% 5/30/2058	-6.79	7.32	0.14	2.70	1.3390	23.4571	0.0087	0.1722
(56) 0% 6/22/2029	-4.13	4.83	-0.01	1.28	1.3346	4.0993	0.2696	2.5954
(57) 0.5% 5/24/2055	-8.51	8.84	0.23	2.17	0.7118	28.9378	0.0339	2.0984
(58) 0.5% 6/28/2045	-7.86	7.94	0.11	1.10	0.5003	0.2811	9.9921	30.7706
(59) 0.5% 6/27/2032	-2.52	2.58	0.05	0.88	0.8578	0.3493	9.0653	41.7957
All yields	-0.22	9.08	0.09	5.78	-	-	-	-

Table 2: **Pricing Errors and Estimated Bond-Specific Risk Parameters**
This table reports the mean pricing errors (Mean) and the root mean-squared pricing errors (RMSE) of Swiss Confederation bonds in the AFNS and AFNS-R models estimated with a diagonal specification of $K^{\mathbb{P}}$ and Σ. All errors are reported in basis points. Also reported are the estimates of the bond-specific parameters within the AFNS-R model and their associated standard errors (SE). Note that standard errors are not available (n.a.) for the normalized value of β^{29}.

C Accuracy of the Swiss Confederation Bond Price Data

In this appendix, we assess the quality of our Swiss Confederation bond data. First, we describe how we construct our own synthetic zero-coupon yields based on the bond price data before we go on to compare them to the zero-coupon yields constructed by SNB staff and made publicly available on the SNB website. Finally, to identify any major differences, we compare the fit of both of these synthetic yield curves to the fit of various dynamic yield curve models estimated using our data.

C.1 Construction of Synthetic Zero-Coupon Yields

We construct synthetic zero-coupon yields using the Svensson (1995) discount function combined with the panel of Swiss Confederation coupon bond prices described in the main text. The Svensson (1995) yield curve has a flexible functional form given by

$$y_t(\tau) = \beta_0(t) + \frac{1 - e^{-\lambda_1 \tau}}{\lambda_1 \tau}\beta_1(t) + \Big(\frac{1 - e^{-\lambda_1 \tau}}{\lambda_1 \tau} - e^{-\lambda_1 \tau}\Big)\beta_2(t) + \Big(\frac{1 - e^{-\lambda_2 \tau}}{\lambda_2 \tau} - e^{-\lambda_2 \tau}\Big)\beta_3(t), \quad (1)$$

where we impose the restrictions that $\lambda_1 > \lambda_2 > 0$. This function contains the level, slope, and curvature components known from Nelson and Siegel (1987) and augments them with an additional curvature factor to provide a better fit to the long end of the yield curve. The corresponding discount function is easily obtained as $P_t^{zc}(\tau) = e^{-y_t(\tau)\tau}$. Now, consider the value at time t of a coupon bond with maturity at $t + \tau$ that pays an annual coupon C. Its clean price, denoted $P_t(\tau, C)$, is simply the sum of its remaining cash flow payments weighted by the zero-coupon bond price function $P_t^{zc}(\tau)$:

$$P_t(\tau, C) = C(t_1 - t)P_t^{zc}(t_1) + \sum_{j=2}^{N} \frac{C}{2} P_t^{zc}(t_j) + P_t^{zc}(\tau), \quad t < t_1 < \ldots < t_N = \tau. \quad (2)$$

We estimate the parameters in the Svensson (1995) curve, $\psi = (\beta_0, \beta_1, \beta_2, \beta_3, \lambda_1, \lambda_2)$, for each observation date by optimizing the following objective function:

$$\min_{\psi} \sum_{i=1}^{n_t} \frac{1}{D_t^{Data,i}} (P_t^{Data,i} - \widehat{P_t^i}(\psi))^2, \quad (3)$$

where n_t is the number of coupon bond prices observed on day t, $P_t^{Data,i}$ is the observed price for bond number i, $\widehat{P_t^i}$ is its price implied by the Svensson (1995) discount function, and $D_t^{Data,i}$ is its duration, which is model-free and calculated before the estimation based on the Macaulay formula. The stated objective is to minimize the weighted sum of the squared deviations between the actual bond prices and the predicted prices, where the weights are the inverse of the durations of each individual security. This is identical to the objective function used by Gürkaynak et al. (2007, 2010). The optimization for each observation date is started at the same parameter vector:

$$\begin{pmatrix} \beta_0 \\ \beta_1 \\ \beta_2 \\ \beta_3 \\ \lambda_1 \\ \lambda_2 \end{pmatrix} = \begin{pmatrix} 0.05010075 \\ -0.02569675 \\ 0.02326028 \\ -0.01846854 \\ 0.8378759 \\ 0.09652915 \end{pmatrix}.$$

The purpose of these synthetic yields is twofold. First, they serve as a key input into the validation of the fit of the dynamic term structure models we consider in the paper. Second, they can be used to evaluate the quality of our Swiss Confederation bond data by comparing our synthetic yields to those produced and made publicly available by the SNB.[1]

C.2 Synthetic Zero-Coupon Yield Comparison

In this section, we focus on the synthetic yields constructed with the Svensson (1995) yield model as described in Appendix C.1. Since this is the same yield model used by SNB staff, we can perform a clean apples-to-apples comparison of the synthetic yields from these two samples. Provided the yields are very close to each other, it would imply that our bond price data are qualitatively very similar to those used by SNB staff.

Table 3 reports summary statistics for the differences between the two data sets at various maturities. The mean absolute differences for yields in the three- to thirty-year maturity range are within 4 basis points and hence small. Larger deviations emerge at the shortest maturities, with mean absolute differences at the one-year maturity of 14 basis points. Large maximum outlier differences are also evident at both ends of the yield curve. Non-negligible discrepancies are also evident in the correlations between the two data sets, shown in the last two columns in Table 3. The correlations are clearly less than one at short and long maturities. On the other hand, it is clear from Table 3 that the constructed yield curves are almost indistinguishable in the range from three to twenty years remaining to maturity.

[1] The parameters from the Svensson (1995) yield curve fitted daily by SNB staff are available at the link:

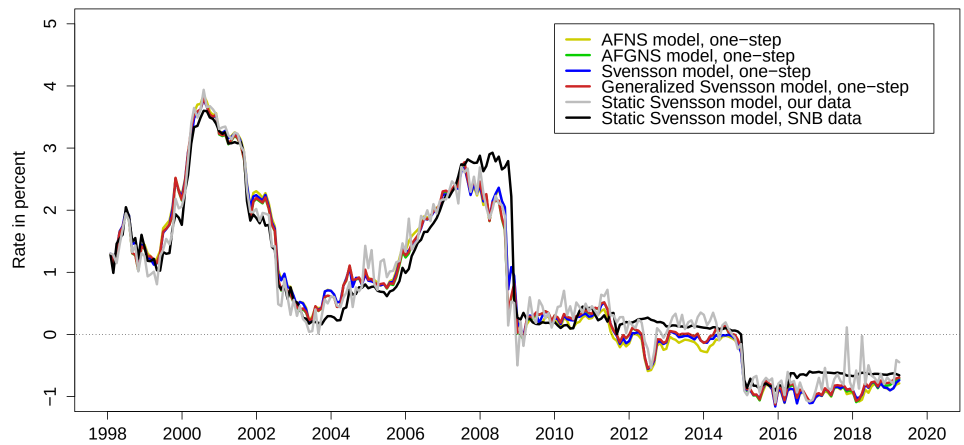
Maturity in years	Mean diff.	Mean abs. diff.	Max. abs. diff.	Correlation	
				Levels	Diff.
1	8.07	13.57	179.27	0.988	0.626
2	5.44	7.46	37.49	0.999	0.927
3	1.27	3.67	18.77	1.000	0.965
4	-0.84	2.47	14.16	1.000	0.978
5	-1.54	2.45	8.68	1.000	0.981
6	-1.56	2.32	9.23	1.000	0.983
7	-1.35	2.09	10.34	1.000	0.984
8	-1.12	1.87	12.50	1.000	0.985
9	-0.92	1.74	15.24	1.000	0.986
10	-0.79	1.74	15.51	1.000	0.986
11	-0.72	1.88	13.55	1.000	0.985
12	-0.69	2.03	14.35	1.000	0.984
13	-0.70	2.16	16.67	1.000	0.981
14	-0.72	2.29	18.31	1.000	0.978
15	-0.76	2.42	20.64	1.000	0.972
16	-0.80	2.52	22.45	1.000	0.964
17	-0.84	2.57	26.16	1.000	0.954
18	-0.88	2.58	34.78	0.999	0.941
19	-0.91	2.56	43.46	0.999	0.926
20	-0.93	2.54	52.11	0.999	0.908
21	-0.94	2.50	60.62	0.999	0.888
22	-0.94	2.45	68.95	0.999	0.866
23	-0.93	2.41	77.05	0.999	0.841
24	-0.90	2.41	84.88	0.999	0.815
25	-0.87	2.46	92.44	0.999	0.787
26	-0.83	2.64	99.70	0.999	0.757
27	-0.77	2.88	106.66	0.998	0.727
28	-0.71	3.16	113.32	0.998	0.695
29	-0.64	3.48	119.69	0.998	0.662
30	-0.56	3.83	125.78	0.998	0.629

Table 3: **Comparing Two Data Sets of Synthetic Zero-Coupon Yields**
The table reports the summary statistics for the mean differences, the mean absolute differences, and the maximum absolute differences between synthetic Swiss zero-coupon yields from the Swiss National Bank and our implementation of the Svensson (1995) discount function. These differences are reported in annual basis points. The last two columns report the correlations between the two yield series for each maturity in levels and first differences, respectively. The data series are monthly, covering the period from January 30, 1998, to March 29, 2019.

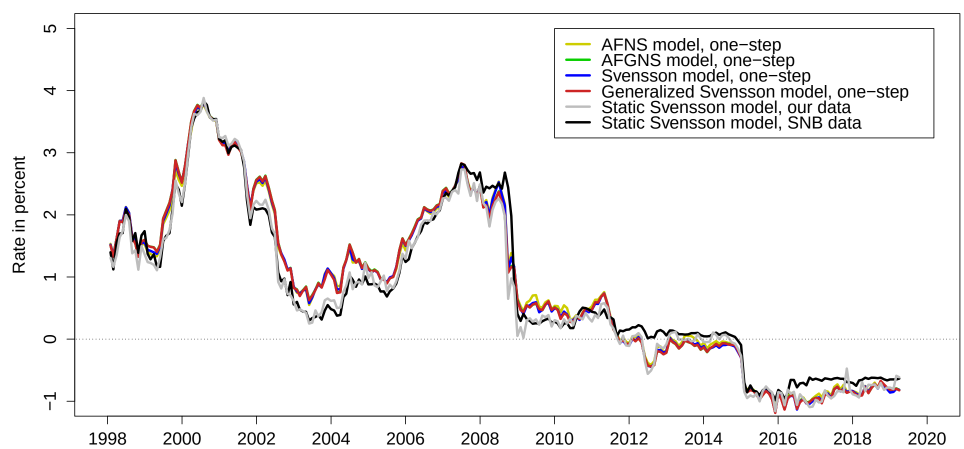
These observations raise the question: which yield curve is the better representation? And is there a better way to construct synthetic yield curves?

To address these questions, Andreasen et al. (2019) recommend using one-step estimations and focusing on the fitted yield curves from such exercises. Therefore, we follow their

https://data.snb.ch/en/topics/ziredev#!/cube/rendopar



(a) One-year yield



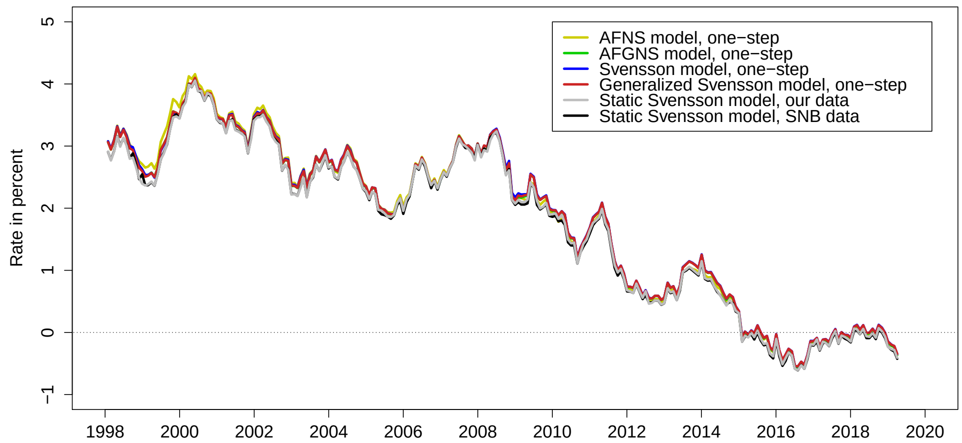
(b) Two-year yield

Figure 1: **Comparison of Synthetic Shorter-Term Zero-Coupon Yields**

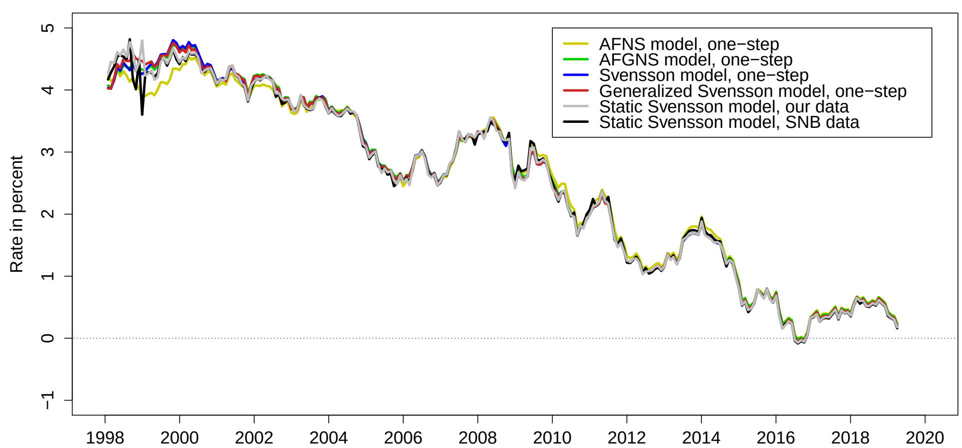
advice and estimate the four dynamic term structure models described in Section 4.4 of the paper with the one-step method. We include the fitted yields from these four models in the comparison.

Figure 1 provides comparisons of the one- and two-year Swiss zero-coupon yields constructed in these six different ways, while Figure 2 shows the corresponding comparisons for the ten- and thirty-year Swiss zero-coupon yields.

We note that short-term yields are sensitive to the yield curve construction method used. We also note that the short-term zero-coupon yields constructed by SNB staff are different for extended periods from the other five yield sets. This suggests that the SNB yields may be



(a) Ten-year yield



(b) Thirty-year yield

Figure 2: **Comparison of Synthetic Longer-Term Zero-Coupon Yields**

constructed using only a subset of our sample of bond prices. Alternatively, it may include some additional short-term interest rate information, which anchors the short end of these curves at a different level than the other curves. Finally, it is notable that the static Svensson model produces yield curves with significantly more volatile short-term interest rates. We ascribe this to the relatively low representation of short-term bonds in our sample for much of our sample period.

On the other hand, interestingly, any such differences are much less pronounced when it comes to the constructed long-term yields as evidenced in Figure 2. Indeed, for the benchmark ten-year yields, there are practically no material differences across any of the six series shown.

Maturity bucket in years	No. obs.	One-step estimation								Static Svensson model			
		AFNS		AFGNS		Svensson		Gen. Svensson		Our data		SNB data	
		Mean	RMSE	Mean	RMSE	Mean	RMSE	Mean	RMSE	Mean	RMSE	Mean	RMSE
0-2	704	-0.39	15.71	-0.09	10.33	-0.59	13.84	-0.30	10.35	-2.26	16.53	-6.69	24.43
2-4	737	0.79	8.44	0.47	5.94	1.40	5.85	1.12	5.98	2.20	5.89	1.08	6.42
4-6	655	-0.84	6.48	-1.00	5.06	-0.74	5.41	-0.99	5.05	-0.92	5.20	0.53	5.46
6-8	570	-0.36	5.37	-0.49	5.40	-0.95	4.96	-1.04	5.40	-1.58	4.85	-0.37	4.98
8-10	546	-0.57	5.85	-0.24	4.08	-1.27	4.14	-0.89	3.98	-1.67	3.78	-0.86	3.71
10-12	461	-0.21	5.65	0.34	3.68	-0.59	3.94	-0.21	3.64	-0.95	3.68	-0.10	3.35
12-14	317	-0.28	6.50	0.81	4.06	0.67	4.19	0.62	3.95	0.62	3.78	0.92	4.59
14-16	212	1.24	7.62	2.00	5.98	2.08	5.41	2.25	5.95	1.87	4.64	2.60	6.05
16-18	157	1.19	5.88	0.78	3.77	1.75	3.74	1.30	3.69	2.23	3.91	2.36	4.75
18-20	183	2.45	7.45	2.06	6.71	2.82	6.83	2.67	6.77	1.93	5.53	4.25	7.57
20-22	120	-0.65	6.08	-0.11	5.59	1.51	5.52	1.19	5.52	1.34	4.71	1.65	5.63
22-24	132	-3.09	7.24	-3.38	5.36	-1.60	4.30	-2.16	4.38	-0.86	3.61	-0.73	3.87
24-26	108	-3.70	8.51	-3.30	7.22	-1.50	6.46	-1.99	6.66	-1.91	5.76	-1.07	6.28
26-28	118	-7.55	9.64	-7.38	8.91	-5.09	6.98	-5.44	7.29	-4.64	6.01	-4.55	6.47
28-30	90	-5.33	10.98	-7.60	9.03	-6.80	8.20	-6.95	8.37	-5.68	7.28	-5.70	7.36
30<	344	1.65	10.14	0.51	6.83	-0.64	6.40	-0.20	6.39	0.09	4.21	-0.09	7.15
All bonds	5,454	-0.31	8.79	-0.30	6.34	-0.28	7.02	-0.28	6.21	-0.50	7.46	-0.62	10.16

Table 4: **Summary Statistics of Fitted Errors of Swiss Confederation Bond Yields**
This table reports the mean pricing errors (Mean) and the root mean-squared pricing errors (RMSE) of the Swiss bond prices for various models estimated on the sample of Swiss Confederation bond prices. The pricing errors are reported in basis points and computed as the difference between the implied yield on the coupon bond and the model-implied yield on this bond. The data are monthly and cover the period from January 30, 1998, to March 29, 2019.

Thus, in the core five- to twenty-year maturity segment, the construction of zero-coupon yields is very robust and barely sensitive to the yield curve construction method used.

C.3 Analysis of Model Fit

To begin, we compare the fit we obtain with our yield curve models to that produced by the synthetic yield curve constructed daily by SNB staff and made publicly available.[2] Presumably, SNB staff would only include high quality bond prices in their yield curve construction. Hence, if those yield curves are able to provide a close fit to our data, the logical implication must be that our data are of high quality. Furthermore, if our models deliver a fit that is on par or better than that obtained with the synthetic yield curves produced by SNB staff, it would be a comforting sign suggesting that our models are flexible and able to deliver what must be characterized as an accurate fit to the prices trading in the Swiss Confederation bond market. As a consequence, their fitted yield curves would be a reasonable representation of historical Swiss Confederation bond yield curves that could be used for both academic research and monetary policy analysis.

[2]Please note that this exercise covers data only back to 1998, when the SNB yield data become available.

In Table 4, we compare the fit of four yield curve models that we implement with a one-step estimation as recommended by Andreasen et al. (2019). Furthermore, we include the fit from our synthetic zero-coupon yield curves as well as those provided by the SNB. First, the yield curves produced by SNB staff deliver a satisfying fit with an overall RMSE of 10.16 basis points. Furthermore, its fit to bond prices with maturities greater than two years is very tight. Second, our own Svensson (1995)-based synthetic yield curves deliver an even better fit, which is not surprising given that it was fitted to this exact data sample. These two observations combined lead us to conclude that our Swiss Confederation bond data are indeed of high quality.

Finally, as for the four models considered in the one-step estimations, we note that there is only very marginal improvement from increasing flexibility by switching from the three-factor AFNS model to the very flexible AFGNS model or the generalized Svensson model, which both have five factors. We take this as further evidence that our augmented AFNS-R model is well-specified regarding its structure for the frictionless level, slope, and curvature factors.

D Description of Regression Variables

In this appendix, we provide a detailed description of the variables used in the regression analysis in the paper, starting with our core set of control variables, followed by details about the additional set of control variables used in the analysis.

D.1 Core Control Variables

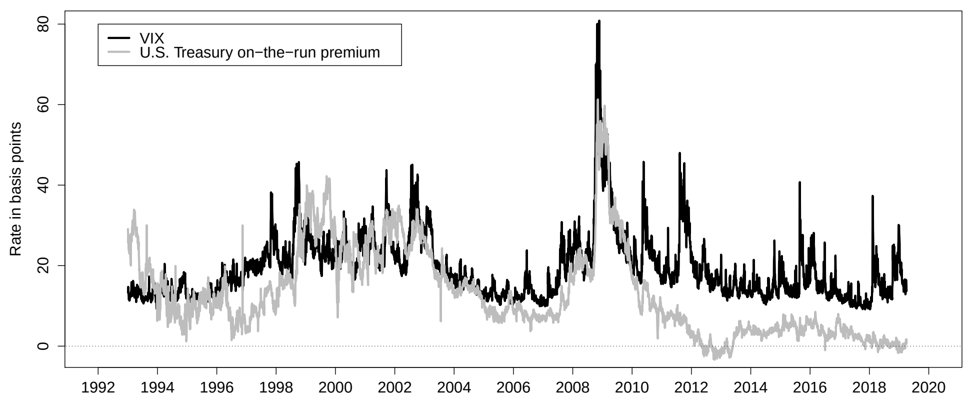


Figure 3: **VIX and U.S. Treasury On-the-Run Premium**

Figure 3 shows the CBOE Volatility Index (VIX) along with the U.S. Treasury ten-year on-the-run premium derived from the difference between par-coupon yields of seasoned ten-year Treasury bonds (as per Gürkaynak et al. (2007)) and the yields on newly issued ten-year Treasury bonds (as reported in the Federal Reserve's H.15 series). We note the high positive correlation (65%) between these two measures of financial market uncertainty and risk aversion.

Figure 4 shows the three-month CHF LIBOR, which has been the leading policy rate of the SNB since 2000, when its current monetary policy framework was adopted. Furthermore, in private conversations, SNB staff confirmed that this rate is also a representative measure of the stance of Swiss monetary policy under the previous monetary policy framework. As a consequence, we use this series as a measure of the opportunity costs of holding money, which has been shown by Nagel (2016) to be a good proxy for the liquidity premium in U.S. Treasury bills, and we want to control for similar effects in the pricing of Swiss Confederation bonds.

The figure also shows the ten-year Swiss zero-coupon yield implied by the AFGNS model

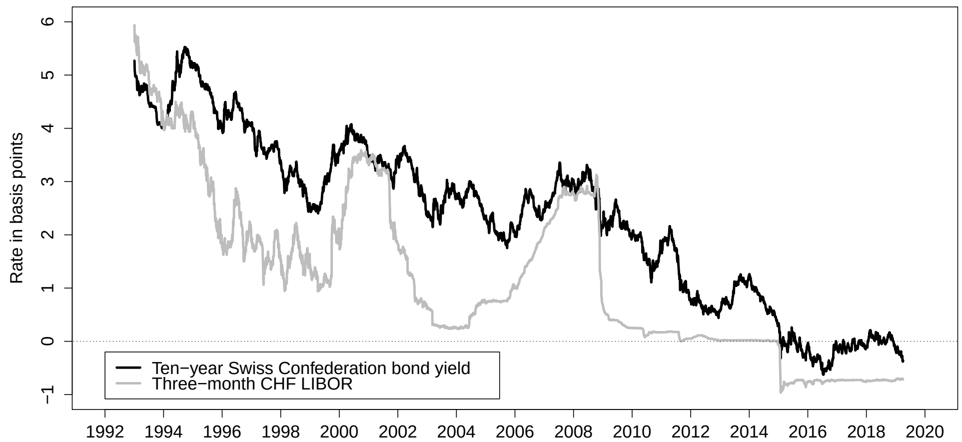


Figure 4: **Swiss Interest Rates**
Illustration of the fitted ten-year Swiss Confederation bond yield implied by the AFGNS model along with the three-month CHF LIBOR rate downloaded from Bloomberg.

considered in Section 4.4 of the paper and estimated with a one-step approach as recommended by Andreasen et al. (2019). This series is used to construct the yield spread between German and Swiss ten-year yields and, as demonstrated in Appendix C.2, it is very robustly estimated.

Figure 5 shows the German and Italian ten-year yields downloaded from Bloomberg. In our regression analysis, we include the yield spread between the two to control for (1) effects tied to the compression of sovereign yield spreads in the run-up to the launch of the euro in January 1999; and (2) flight-to-safety effects during various phases of the European sovereign debt crisis. Furthermore, we use the German ten-year yield to calculate the yield spread relative to the matching Swiss ten-year yield shown in Figure 4. Note also that the convergence between the German and Italian yields starts well ahead of the euro launch, but is not complete until relatively close to January 1999, which is a time series pattern that is largely consistent with the time profile of the level shift we observe in the Swiss safety premium around the time of the launch of the euro.

The next variable is the U.S. TED spread, which is calculated as the difference between the three-month U.S. LIBOR and the three-month U.S. T-bill interest rate and shown in Figure 6. This spread represents a measure of the perceived general credit risk in global financial markets that could affect the pricing and trading of Swiss Confederation bonds.

Figure 7 shows the Swiss public debt-to-GDP ratio, which serves as an important control for changes in the supply of Swiss Confederation bonds and as a proxy for the true, but unob-

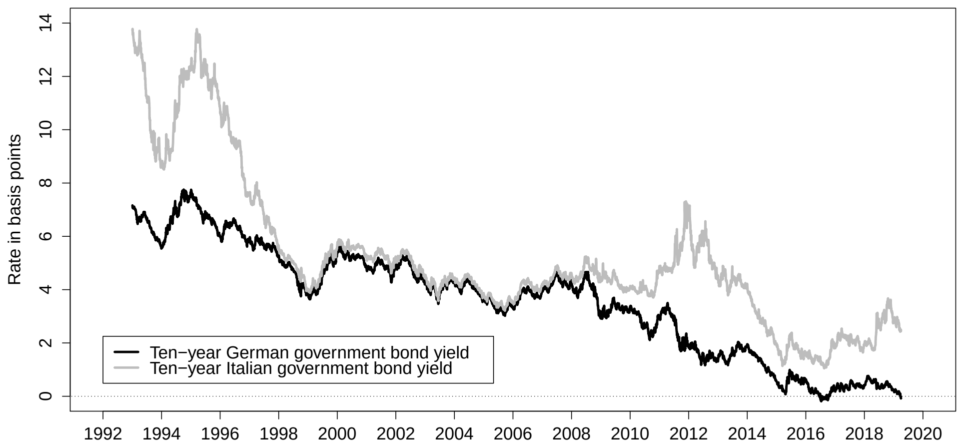


Figure 5: **German and Italian Ten-Year Government Yields**
Illustration of ten-year German and Italian government bond yields.

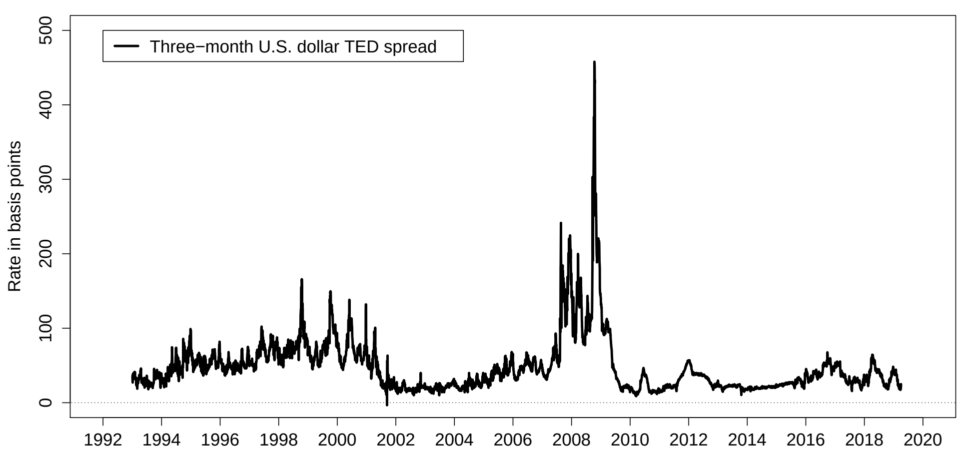


Figure 6: **Three-Month U.S. TED Spread**

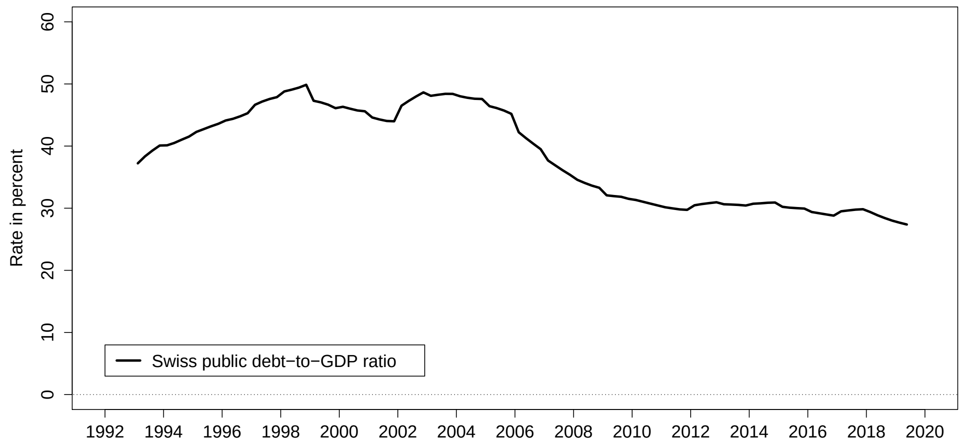


Figure 7: **Swiss Public Debt-to-GDP Ratio**
Illustration of the Swiss public debt-to-GDP ratio. The data series is quarterly and therefore linearly interpolated to produce the monthly series used in the regression analysis.

served variation in the stock of Swiss safe assets following the analysis of Krishnamurthy and Vissing-Jorgensen (2012). Note that this series is quarterly. Hence, we use linear interpolation to convert it into a monthly series that can be used in the regression analysis.

Following the work of Houweling et al. (2005), we include the average Confederation bond age as a proxy for bond liquidity. This series is shown in Figure 8 along with the average remaining time to maturity. We note that both the age and the duration of the available universe of Swiss Confederation bonds have trended up during the sample period and are near all-time highs at the end of our sample with a value of 8.68 years and 15.55 years, respectively.

Houweling et al. (2005) also recommend using yield volatility as a proxy for bond liquidity. To that end, we use a standard measure of realized volatility based on daily data. First, we estimate the AFGNS model using the daily sample of Swiss Confederation bond prices considered in Section 5.2 of the paper. This gives us daily fitted Swiss Confederation zero-coupon yields at all relevant maturities. We then generate the realized standard deviation of daily changes in interest rates for the past 31-day period on a rolling basis. The realized variance measure is used by Andersen and Benzoni (2010), Collin-Dufresne et al. (2009), and Jacobs and Karoui (2009) in their assessments of stochastic volatility models. This measure is fully nonparametric and has been shown to converge to the underlying realization of the conditional variance as the sampling frequency increases; see Andersen et al. (2003) for

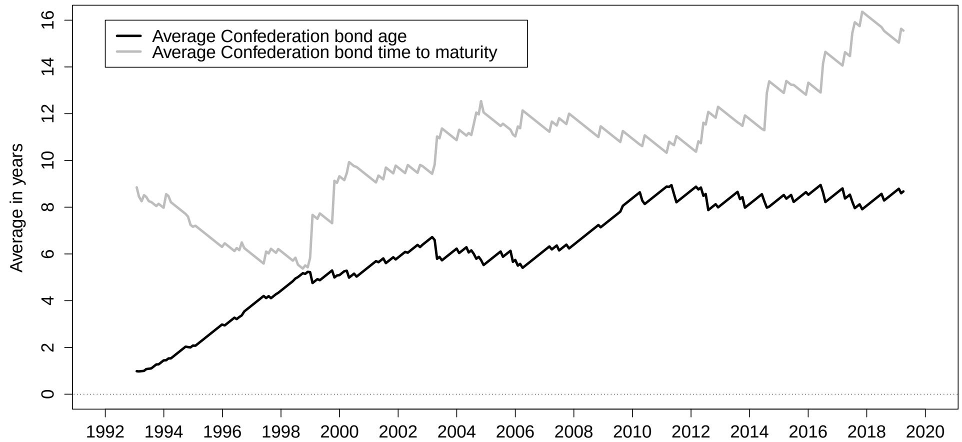


Figure 8: **Average Confederation Bond Age and Time to Maturity**
Illustration of the average age and time to maturity of the Swiss Confederation bonds included in the sample, which covers the period from January 29, 1993 to March 29, 2019 and censors each bond's price when it has less than three months to maturity.

details. The square root of this measure retains these properties. For each observation date t we determine the number of trading days N during the past 31-day time window (where N is most often 21 or 22).[3] We then generate the realized standard deviation as

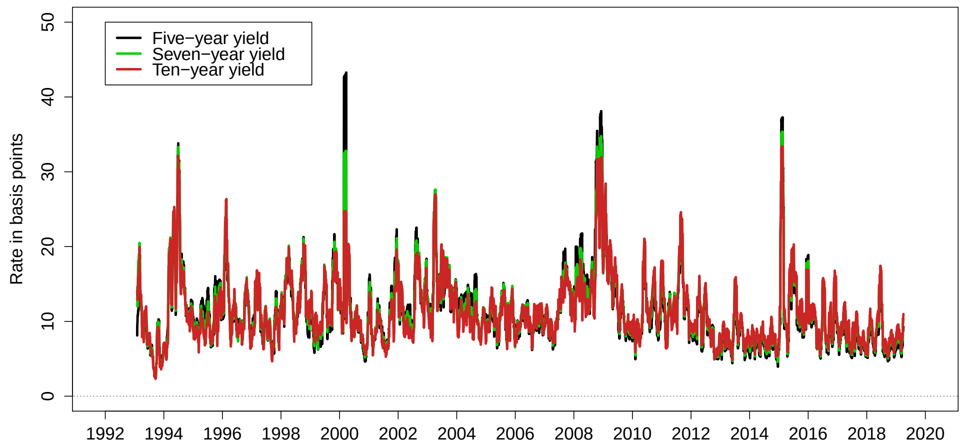
$$RV_{t,\tau}^{STD} = \sqrt{\sum_{n=1}^{N} \Delta y_{t+n/N}^2(\tau)},$$

where $\Delta y_{t+n/N}(\tau)$ is the change in yield $y(\tau)$ from trading day $(n-1)$ to trading day n.[4]

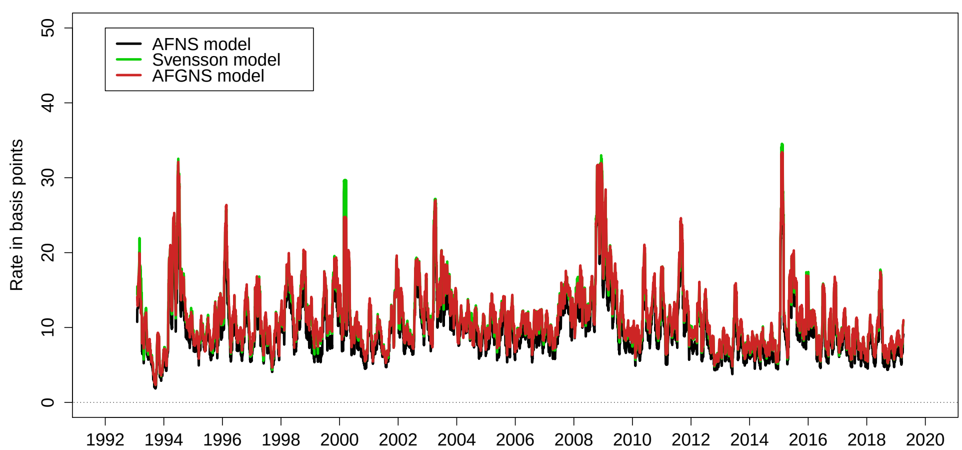
Figure 9 shows that the realized yield volatility series used in the regression analysis is not sensitive to the choice of the yield maturity or the model used to construct the fitted zero-coupon yields. To achieve the greatest accuracy in the constructed yields, we follow Andreasen et al. (2019) and focus on the fitted yield implied by the five-factor AFGNS model. Furthermore, given that the average time to maturity of the available Confederation bonds is close to ten years for much of our sample period as demonstrated in Figure 8, we choose to use the one-month realized volatility of the ten-year yield in our regressions, but we stress that our results are clearly robust to alternative choices, both in terms of the model used and

[3]As a consequence, the realized volatility measure can be calculated for the period from February 4, 1993, to March 29, 2019.

[4]Note that other measures of realized volatility have been used in the literature, such as the realized mean absolute deviation measure as well as fitted GARCH estimates. Collin-Dufresne et al. (2009) also use option-implied volatility as a measure of realized volatility.



(a) Comparison across maturities



(b) Comparison across models

Figure 9: **Realized Yield Volatility Comparisons**
The top panel shows the one-month realized volatility of yields with three different maturities, all constructed from the AFGNS model estimated with our daily sample of Swiss Confederation bond prices. The bottom panel shows the one-month realized volatility of Swiss ten-year yields constructed from three different yield curve models, all estimated with our daily sample of Swiss Confederation bond prices.

the maturity considered.

Inspired by the analysis of Hu et al. (2013), we also include a noise measure of Swiss Confederation bond prices to control for variation in the amount of arbitrage capital available in this market. In principle, this could be constructed using any yield curve model. However, to be consistent with the rest of the analysis and limit focus to models that have been used in the existing literature, we choose to focus on the Svensson (1995) model and the AFGNS

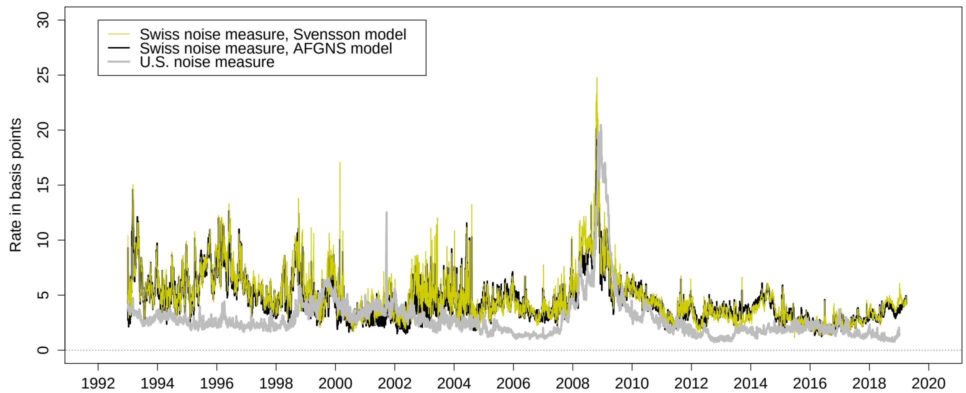


Figure 10: **Noise Measures**
Illustration of the Swiss noise measures constructed based on the fitted errors implied by the estimated Svensson and AFGNS models using the Swiss Confederation bonds covering the period from January 4, 1993, to March 29, 2019, with a comparison to the U.S. noise measure constructed by Hu et al. (2013).

model already used elsewhere, both estimated with the one-step approach recommended by Andreasen et al. (2019). The two resulting noise measures, defined as the average absolute fitted errors of the yield to maturity across all available bonds at each observation date, are shown in Figure 10 with a comparison to the U.S. noise measure constructed by Hu et al. (2013) and accessed from Jun Pan's personal website with data through the end of 2018.[5] The correlation between the noise measure constructed based on the estimated Svensson model and that based on the estimated AFGNS model is 94.8 percent at daily frequency. Thus, this measure is not very sensitive to the specific yield curve model used, as also emphasized by Hu et al. (2013).

D.2 Additional Control Variables

Besides the set of core control variables, we consider several additional confounding factors in the regressions. We add the overnight federal funds rate shown in Figure 11 to proxy for the U.S. safe-asset liquidity premium as in Nagel (2016), and reported earnings per share of companies in the S&P 500 to account for opportunity costs in the equity market. We also consider the MOVE volatility index to proxy for risk aversion in the bond market. Finally, we include the total sight deposits at the SNB to control for any possible reserve-induced effects of the SNB's FX interventions; see Christensen and Krogstrup (2019).

[5]See the link: http://en.saif.sjtu.edu.cn/junpan/

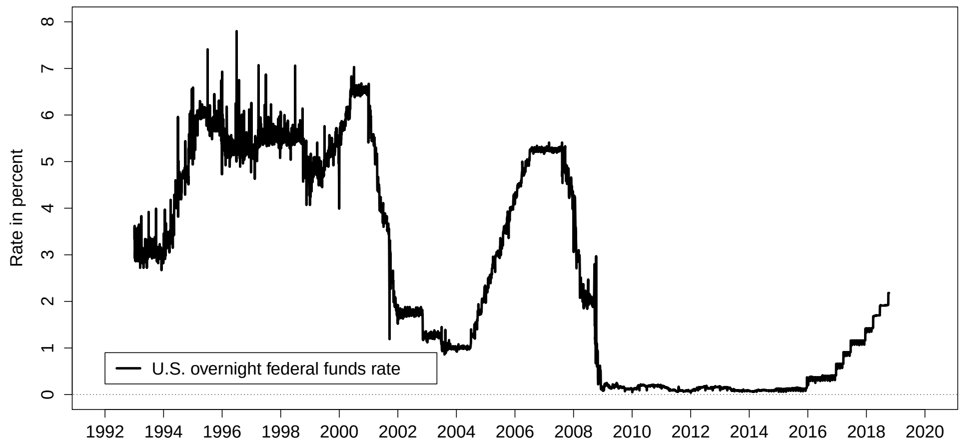


Figure 11: **U.S. Overnight Federal Funds Rate**

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