Comparative Analysis of Risk in Euro-Area Sovereign Yields

Sophia Azimi

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Abstract
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Keywords
Euro, EMU, sovereign bonds, risk premia, Kalman filtering

Disciplines
Business

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COMPARATIVE ANALYSIS OF RISK PREMIA IN EURO-AREA SOVEREIGN YIELDS

By

Sophia Azimi

An Undergraduate Thesis submitted in partial fulfillment of the requirements for the

JOSEPH WHARTON SCHOLARS

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APRIL 2022
ABSTRACT

What risk premia exist in Euro-area sovereign yields, and how big are they? What are the similarities and differences across countries? This paper aims to answer these questions for eight Euro countries by using Kalman filtering to decompose their sovereign yields into the following premia: a short-rate and term premium, a default premium, a liquidity premium, and a segmentation premium. The main finding is that countries with similar credit ratings and yield magnitudes tend to exhibit similar patterns, and that on average, countries that were especially affected by the European Debt Crisis (Italy, Portugal, Spain) tend to have default premia that account for more than 50% of their sovereign yields.

Keywords

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INTRODUCTION

The Euro: An Economic Experiment

Has the creation of the Euro been a success? Since the European Monetary Union was first founded in 1999, a total of nineteen European countries\(^1\) have joined. The goal of creating this common currency was to make cross-border trade easier, the European economy more stable, and to provide consumers with a greater set of economic opportunities.\(^2\) However, to what extent have these economic goals been achieved?

Moreover, beyond its economics, the Euro is a topic of social and political debate. Both the Global Financial Crisis (GFC) and European Debt Crisis (EDC) made the currency increasingly unpopular. In this period of crisis, trust in the European Central Bank (ECB) decreased from 30% in 2008 to a record low of below -20% in 2014, and net support for the Euro reached a record low of 30% in 2013.\(^3\) Populist movements also began to gain more traction, especially as these two economic crises led to increased unemployment and social dissatisfaction.\(^4\) The currency and its success, therefore, affect a broad range of stakeholders, not only within the EMU, but also beyond it as the Euro remains an important global currency.

Evaluating the success of the Euro, therefore, is a question of great interest for its various stakeholders. The methods one can use to measure economic progress are broad. Some past literature involves measuring a given country’s macroeconomic variables prior and subsequent to adoption of the Euro. Other papers seek to measure convergence of the various Euro sovereigns by comparing macroeconomic variables across countries. In the following sections, I seek to

\(^1\) Austria, Belgium, Cyprus, Estonia, Finland, France, Germany, Greece, Ireland, Italy, Latvia, Lithuania, Luxembourg, Malta, Netherlands, Portugal, Slovakia, Slovenia, and Spain

\(^2\) European Union.

\(^3\) Bergbauer, S., N. Hernborg, J. Jamet, E. Persson, and H. Schölermann.

\(^4\) Algan, Y., S. Guriev, E. Papaioannou, and E. Passari.
provide an overview of the literature thus far relating to the measurement and comparison of Euro-area countries across different financial indicators.

**Literature Review**

**Early (Pre-GFC) Evaluations of the EMU**

First, a couple of early papers from Lane (2006) and Wyplosz (2006) sought to characterize the initial progress of the EMU, notably in the years preceding the GFC.

Lane’s 2006 paper provided some basic summary statistics by country of different economic indicators, measuring their change over a given time period. However, these statistics did not account for any variation or underlying heterogeneity among the countries, but rather demonstrated simple increases or decreases. Some indicators measured included the change in average annual inflation rates, the evolution of national competitiveness (as measured by changes in relative price levels over the period), the proportion of a country’s international portfolio holdings allocated to other EMU countries, trade/GDP ratio, and trade share with other EMU countries. Key takeaways included that there have been persistent differences in national inflation rates since the common monetary policy has not been one-size-fits-all, as well as that the EMU has contributed to greater cross-border trade in finance and goods due to higher market integration.

As the name of his paper, "European Monetary Union: the dark sides of a major success" suggests, Wyplosz (2006) took a more pessimistic perspective on the creation of the union. This predominantly qualitative paper mainly looked at flaws in the founding of the EMU and how economic theory that was laid out as the rationale for building the currency union has not always been followed, leading to suboptimal outcomes.
Comparing Macroeconomic Indicators Across EU/EMU Countries

**General comparisons.** A recent paper by Gehringer and Konig (2021) looked at the degree of economic integration and business cycle synchronicity among EU member countries (inclusive of the Euro countries) through an analysis of several macroeconomic variables. They classified their metrics to measure integration under two categories: nominal and real. Nominal integration variables included adjusted GDP growth rate, inflation rate, core inflation rate, change in employment, and balance of government budget as a percentage of GDP, while real integration variables included the overall unemployment rate, youth unemployment, labor productivity, and the industrial production growth rate.

Moreover, as opposed to reporting by country, Gehringer and Konig reported results by the following groups: EU, Euro, Euro Core\(^5\), Euro Periphery\(^6\), and EU Non-Euro\(^7\). Their methodology was fairly simple; the authors calculated pairwise Pearson correlation coefficients for the aforementioned macroeconomic indicators across all pairings of countries within each group, then averaged these to obtain the average correlation for EU, Euro, Euro Core, Euro Periphery, and EU Non-Euro. They then repeated this process for different time periods. Overall, they found that integration up-to-date has not been successful; most of the indicators showed declining and low correlation coefficients across member states.

A 2018 working paper by Franks et al. also performed a general analysis of economic convergence in the Euro area. They found that there has been nominal convergence of inflation and interest rates but real convergence of per capita income levels has not occurred since the

---

\(^5\) Austria, Belgium, Finland, France, Germany, and the Netherlands

\(^6\) Greece, Ireland, Italy, Portugal and Spain

\(^7\) Bulgaria, Croatia, Czech Republic, Denmark, Hungary, Poland, Romania, and Sweden
common adoption of the Euro. Looking at inflation rates between 1999-2007, Franks et al. believe that consistent inflation differentials led to deterioration of Ireland, Greece, Spain and Portugal's competitiveness over time, despite overall variation in inflation rates being fairly small.

**Business Cycles.** In a 2020 paper by Duarte and Gehringer, the authors looked at the assumptions behind the synchronization of business cycles and eventual income convergence. They presented a hypothetical mathematical argument, comparing a poor country to a rich one, to demonstrate that synchronized business cycles and correlated growth rates do not necessarily guarantee real income convergence, in line with the Optimum Currency Area theory that helped pioneer the foundation of the EMU. While the paper does not make any use of empirical evidence from the actual countries within the EMU, they present their argument as generalizable to the union.

**Productivity.** A paper by Papaioannou (2021) used the synthetic control method for causal inference to estimate whether entering the EMU affected Total Factor Productivity (TFP) of the twelve founding countries of the EMU. The author does acknowledge, however, that the link between integration and productivity is a subject of theoretical debate (i.e. higher productivity may not necessarily be a required condition for convergence to occur). Results demonstrated that post-Euro adoption, countries' TFPs improved, with the exception of Portugal. Greece and Italy also demonstrated minimal improvement.

Looking into the econometric methodology used, Papaioannou compared the actual outcome of a country $i$ at time $t$ to the counterfactual outcome that could be observed in a hypothetical country in the absence of the event (i.e. in this case, joining the EMU). Note that

---

8 Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, the Netherlands, Portugal, and Spain
this methodology entailed comparing countries to their progress against themselves as opposed to their convergence and progress relative to other countries in the EMU. In Papaioannou (2021), the synthetic control group consisted of European, Mediterranean, and OECD countries that were not members of the European Union. The author also used a difference in differences estimation on the pre-treatment fit between the actual vs. counterfactual (synthetic) series of productivity outcomes to validate the statistical significance of the comparison between the two for the post-treatment case. A final important methodological decision was to perform the analysis between 1980-2017, hence including the Global Financial Crisis.

**Income Per Capita.** A paper by Puzzello and Gomis-Porqueras (2018) also used the synthetic control method, albeit to estimate outcomes in income per capita for only six EMU countries: Belgium, France, Germany, Ireland, Italy, and the Netherlands. Moreover, the analysis is restricted to 1971-2007, not including the Global Financial Crisis due to the shocks it created. The control group for the synthetic country comparison differs from that of Papaioannou (2021); Puzzello and Gomis-Porqueras (2018) selected the synthetic group of countries based on similar levels of income per capita in the pre-treatment period. However, they also included two extra control groups common to all EMU members: the first being industrial countries as defined by the IMF in 1998 (Australia, Canada, New Zealand, Norway, Singapore, Switzerland, and the USA), and the second being countries with available data and no profound structural shocks during the sample period. Findings demonstrated that income per capita for Belgium, France, Germany, and Italy would have been higher without the adoption of the Euro, that of Ireland would have been lower, and that of the Netherlands would have more or less been the same.

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9 Argentina, Australia, Brazil, Canada, Chile, China, Colombia, Egypt, Hong Kong, Iceland, Indonesia, Israel, Japan, Korea, Malaysia, Mexico, Morocco, New Zealand, Philippines, Switzerland, Thailand, Tunisia, Turkey, and Uruguay
Moreover, Puzello and Gomis Porqueras (2018) also conducted a secondary analysis of the aggregate effects of the Euro through a multiple regression analysis, where the dependent variable was the synthetic estimate of the income per capita gap for the six treatment countries as a percentage of their income per capita observed at a given time (positive values of the dependent variable implied the EMU member in question gained from the Euro while negative values implied the opposite). The main takeaways from this second portion of the paper were that greater benefits from the Euro arose if the country was a) an earlier adopter of the currency with a more synchronized business cycle to that of the union and b) more open to intra-union trade and migration. Integration of capital markets was also an important factor in minimizing post-Euro income losses of a given member of the union.

**Research on the Interaction of Monetary and Fiscal Policy in the EMU**

Another source of relevant research was the interaction between fiscal and monetary policy. Many academics and policy-makers believe that strategic fiscal policy must be implemented in order for the benefits of the currency union to be realized to their greatest potential.

A mostly qualitative recent working paper by Debrun et al. (2021), published by the ECB, provided an overview of the interaction between monetary and fiscal policy in light of the recent economic crises (i.e. the Global Financial Crisis, COVID-19). The authors found that the GFC was highly detrimental to the convergence of the EMU countries, notably as it helped to further expose the heterogeneity of the member countries. They noted that the gap in income per capita for some countries relative to the most productive Member States (such as Germany) had widened further.
Debrun et al. also looked into the heterogeneity of fiscal policy in the EMU, notably by analyzing the size of automatic fiscal stabilizers per country. The authors found that overall, Belgium, France, and Finland ranked the highest, while Latvia, Germany, and Slovakia ranked the lowest. Moreover, another metric the authors looked at was the fall in public investment after the Global Financial Crisis, a) relative to deficits above 3% pre-crisis and b) related to debt increases during the crisis. Countries who exhibited significant decreases in public investment in both charts included Greece, Portugal, Ireland, Spain, Cyprus, and Luxembourg, while countries such as Germany, France, Belgium, Latvia and Estonia increased their public investment.

The authors projected that the debt paths of member countries would continue to show sizeable differences from now until 2030, not only due to COVID-19, but also due to the initial asymmetries from their fundamental economic heterogeneity.

Work done by Andres, Burriel and Shen (2020) also examined the fiscal policy implications of the EMU. In this paper, the authors created a dynamic stochastic general equilibrium model (DSGE) of a two-country union with different levels of respective debt-to-GDP ratios to examine the response of risk premium to public debt levels. In this model, debt is determined endogenously (i.e. by the country itself, as in real-life). The hypothetical high and low debt-to-GDP ratio countries were modeled after Spain and Germany, respectively. They found that the fiscal consolidation necessary to bring a high-debt country's debt ratio down had costs of output loss not only for the country itself, but also for the other members of the union. The key takeaway was that balancing policy moves involves complex dynamics and depends on the macroeconomic environment, notably the prevailing interest rate policy.

**Fiscal and Monetary Policy and Sovereign Bond Yields.** Work by Corradin, Grimm, and Schwaab (2021) contributed to research on the efficacy of Euro area policymaking by
analyzing the impact of EU fiscal policy and EMU monetary policy announcements on the sovereign bond yields of the EMU’s four largest countries by GDP\textsuperscript{10}. They broke yields down into the following components: the expected future short-term risk-free rate and a term premium, a default risk premium, a redenomination risk premium, a liquidity risk premium, and a segmentation (convenience) premium. They found that the importance of different premia depended on a given country and time. For instance, default and redenomination risk premia explained variation for Italian and Spanish bond yields more than French and German ones, while the latter were more driven by expectations on the future short-term risk-free rate, the term premium, and the segmentation premium.

Another interesting finding from this paper was the asymmetric effect of the ECB’s Pandemic Emergency Purchase Programme (PEPP) announcement in March 2020. In response to the PEPP, Italian yields fell significantly, Spanish yields fell moderately, and French and German yields increased. The authors believe this variation occurred due to the PEPP’s flexible nature.

Following Krishnamurthy, Nagel, and Vissing-Jorgensen (2018), Corradin et al. employ Kalman filtering, a statistical method to estimate unobserved variables from observed ones. This technique, as applied to decomposing sovereign yields, is quite novel, and the two aforementioned papers are the only ones to date that use it to estimate the risk premia of Euro-area sovereign yields.

\textsuperscript{10} Germany, France, Italy, and Spain
Purpose and Layout of Study

This research paper is an attempt to add to the literature on sovereign yield decomposition using Kalman filtering. For daily observations between January 2014 - December 2021, I implement the Kalman filtering methodology to decompose the 10-year sovereign yields of eight Euro-area countries into the following premia: a short rate and term premium, a default risk premium, a liquidity premium, and a residual segmentation premium.

There are several advantages of Kalman filtering. First, it involves dynamic linear modeling, which allows for measurements of the latent premia both over time and relative to the sovereign yield (see Methodology section). The Kalman filter’s recursive nature takes into account past-present dependencies which is precisely what allows for the measurement of latent premia over time. Moreover, it is well-suited to make estimations when measurements are noisy, which is an important feature to account for in financial data. Finally, it is easy and computationally-efficient to implement.

This paper will proceed with a Data section, a Methodology section, a Results section, and finally, a Discussion section.

DATA

Selection of Countries

My initial intention was to include all Euro countries in my sample set, where the availability of data permitted. However, upon aggregating the data available on Bloomberg for different countries’ sovereign yields and CDS spreads over January 2014 – December 2021, I narrowed down my sample to the following eight countries, which had the most complete sets of

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11 ScienceDirect.
data: Austria, Belgium, Finland, Italy, Netherlands, Portugal, Slovenia, and Spain. While Krishnamurthy et al. and Corradin et al. looked at Italy, Portugal, and Spain, there is no literature thus far that applies the Kalman filter sovereign yield decomposition approach to Austria, Belgium, Finland, the Netherlands, and Slovenia, providing an opportunity for a novel contribution to the existing literature.

**Selection of Risk Premia**

My initial intention was to measure the same risk premia as the Corradin paper. However, after doing a preliminary search of the Euro data available on Bloomberg, I realized that I had trouble finding measures of redenomination risk. Corradin et al. measured redenomination by using the difference between sovereign CDS spreads under the International Swaps and Derivatives Association (ISDA) contract terms of 2003 as compared to 2014, data which was extremely limited on Bloomberg. While other papers did have alternate methods of measuring redenomination risk, these data too were unavailable on Bloomberg. For example, De Santis (2015) proposed to create a “quanto” CDS measure, which gave the difference in EUR versus USD CDS quotes. A given Euro-area sovereign’s “quanto” CDS would then be compared to another benchmark sovereign’s, the latter generally being Germany. I unfortunately was unable to find any EUR CDS quotes on Bloomberg to replicate this methodology.

Due to the unavailability of data that could be used to measure redenomination risk premia, I therefore decided to focus my analysis on only four latent risk premia: the expected future average short-rate and term premium, the default risk premium, the liquidity risk premium, and the segmentation premium.

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12 Expected future short-term risk-free rate and term premium, default risk premium, redenomination risk premium, liquidity risk premium, and segmentation (convenience) premium
Description and Operationalization of Selected Risk Premia

Short-Rate and Term Premium

The short-rate and term premium is intended to capture the risk from holding the sovereign yield for ten years, including both expectations regarding the future short-term, risk-free interest rate, as well as the risk from holding the bond for an extended period of time into the future (i.e. in this case, ten years). Following both Krishnamurthy et al. and Corradin et al., I use the EONIA overnight index swap rate as a proxy for the short-rate risk and term risk combined.

Default Premium

The default risk premium is intended to capture the extra return investors require to compensate for potential default risk. A common measure of this in previous literature is through analyzing a given sovereign’s CDS spreads (e.g. Rodriguez, Dandapani, and Lawrence 2019, Corradin et al. 2021). Following the existing literature, I also estimate the default risk using CDS spreads.

Moreover, as Krishnamurthy et al. discuss, using USD-quoted sovereign bonds is preferred over EUR-quoted ones because they are better at capturing default risk. In the case that the ECB decided to potentially intervene to keep CDS spreads at a certain rate, this intervention would lower the perception of default risk without truly lowering the actual risk itself. For this reason, I also opt to use CDS spreads quoted in USD as opposed to EUR.

Liquidity Premium

The liquidity premium is intended to account for how easy it is to buy or sell a given sovereign’s debt. Krishnamurthy et al. do not estimate this premium, while Corradin et al. do and find that it is a minimal component of any given sovereign’s yield (they find that it is, on average, between only 3%-6% for France, Germany, Italy, and Spain).
A common, albeit imperfect, method used in past literature to estimate the liquidity premium has been the KfW-Bund spread (Ejsing, Grothe, and Grothe 2012, Schuster and Uhrig-Homburg 2012, Paret and Weber 2019, Corradin et al. 2021). Kreditanstalt für Wiederaufbau (KfW) is a German state-owned investment and development bank whose bonds are guaranteed by the German government. This sovereign guarantee therefore implies that its bonds should have the same default risk as the German government. Furthermore, since KfW bonds are less liquid than German sovereign bonds, the spread between the two can approximate the liquidity premium in Euro bond markets. I use the KfW-Bund spread in this paper to measure the liquidity premium, although it has several limitations which I elaborate on in the Discussion section.

**Segmentation Premium**

The segmentation premium is intended to capture investors’ perceptions of the potential non-pecuniary benefits of holding sovereign bonds. As Krishnamurthy et al. define it, “[i]f investors differ in their valuation of a bond and some investors are constrained from participating in the market, either because they are not active in the market for that bond or because they face portfolio constraints, then the market price will reflect the valuation of a subset of the investor population and bond yields will reflect a segmentation factor compared with the frictionless case” (Krishnamurthy et al., 13).

For instance, one potential benefit would be if the bonds received special regulatory treatment. In the case where investors value the bond beyond its financial value, the segmentation premium would be negative, while when participating investors value the bond less than its cash flows, the segmentation premium would be positive.

Corradin et al. calculate the segmentation premium as a residual value in the sovereign yield and do not use any observed data to directly measure it.
choice. However, since I do not estimate a redenomination premium, it is likely that the residual segmentation premium I estimate may also capture some of the value of the redenomination premium.

**Selection of Bond Maturity and Years Analyzed**

I decide to focus on ten-year sovereign yields because data on ten-year maturities are most readily available on Bloomberg, not only for sovereign yields but also CDS spreads. This is in contrast to Corradin et al. who use five-year maturities for their data and Krishnamurthy et al. who analyze a variety of maturities in their aggregate analysis, ranging from six months to ten years.

With respect to my years of analysis, I decided to begin in 2014 to allow for a couple of years past what could be considered the peak of the European Debt Crisis. I pick an endpoint of December 2021 to include the impact of the COVID-19 pandemic, seeing as it is a current event that severely impacted global financial markets.

**Data Collection**

I used Bloomberg to collect my raw data and tried to match the tickers I used to those of previous Euro-area sovereign research to the extent possible, mainly referencing the tickers in Paret and Weber (2019). Please see Appendix A for a complete list of the tickers I used.

First, for the sovereign yields, I used the ticker for generic bonds of sovereigns, seeing as these track the YTM of a changing benchmark bond over time, and the ten-year sovereign yield of ten years ago will be different from the ten-year sovereign yield of the present. Then, for the EONIA rate, I used the ten-year EONIA OIS rate. For the CDS spreads, I used the USD-denominated ten-year CDS quotes for a given sovereign. Finally, for the KfW-Bund spread, I followed the KfW ticker given in Paret and Weber (2019) to obtain the daily yield for an index.
of ten-year maturity KfW bonds. I then calculated the spread between this aforementioned KfW yield and that of a generic 10-year German sovereign bond using Bloomberg’s Spread Analysis feature to obtain my KfW-Bund spread estimates.

Moreover, to make sure that the data series I used were consistent with the Corradin et al. paper, I looked at the five-year maturity of the same tickers and compared the graphs of the five-year maturity time series to the plots from the Corradin et al. paper, adjusting the years of analysis to match their paper’s (i.e. Corradin et al. used a time series window of January 2015 through October 2020). I found consistency across all my selected tickers and their graphical outputs as compared to the Corradin et al. paper, implying my sources of data were accurate.
Table 1: Summary Statistics for Input Variables\(^{13}\)

<table>
<thead>
<tr>
<th></th>
<th>EONIA Rate</th>
<th>KfW-Bund Spread</th>
<th>Sovereign Yield</th>
<th>CDS Spread</th>
</tr>
</thead>
<tbody>
<tr>
<td>All Countries</td>
<td>0.398</td>
<td>0.238</td>
<td></td>
<td></td>
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<tr>
<td></td>
<td>(0.481)</td>
<td>(0.080)</td>
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<td></td>
</tr>
<tr>
<td>Austria (n = 2044)</td>
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<td></td>
<td>0.434</td>
<td></td>
<td>41.026</td>
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<tr>
<td></td>
<td>(0.567)</td>
<td></td>
<td>(16.630)</td>
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<tr>
<td>Belgium (n = 2044)</td>
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<tr>
<td></td>
<td>0.566</td>
<td></td>
<td>56.438</td>
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<td></td>
<td>(0.617)</td>
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<td>(24.600)</td>
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<td>Finland (n = 2044)</td>
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<tr>
<td></td>
<td>0.398</td>
<td></td>
<td>39.052</td>
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<td></td>
<td>(0.548)</td>
<td></td>
<td>(12.139)</td>
<td></td>
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<td>Italy (n = 2044)</td>
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<td></td>
<td>1.810</td>
<td></td>
<td>193.827</td>
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<td></td>
<td>(0.793)</td>
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<td>(44.687)</td>
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<td>Netherlands (n = 2045)</td>
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<td>0.365</td>
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<td>38.306</td>
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<tr>
<td></td>
<td>(0.587)</td>
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<td>(14.307)</td>
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<td>Portugal (n = 2044)</td>
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<td></td>
<td>1.949</td>
<td></td>
<td>191.805</td>
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<tr>
<td></td>
<td>(1.339)</td>
<td></td>
<td>(105.216)</td>
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<td>Slovenia (n = 2043)</td>
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<tr>
<td></td>
<td>1.074</td>
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<td>144.418</td>
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<tr>
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<td>(1.021)</td>
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<td>(33.071)</td>
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<td>Spain (n = 2043)</td>
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<td></td>
<td>1.272</td>
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<td>114.265</td>
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<tr>
<td></td>
<td>(0.812)</td>
<td></td>
<td>(34.657)</td>
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</tr>
</tbody>
</table>

\(^{13}\) EONIA Rate, KfW-Bund Spread, and Sovereign Yield are quoted in percentage points, CDS spread is quoted in bps. Standard deviations are in parentheses.
**METHODOLOGY**

**Kalman Filter**

Following the work of Corradin et al. and Krishnamurthy et al., I decomposed sovereign yields using Kalman filtering. I implement the Kalman filter using the Fast Kalman Filter ("FKF") package in R.

While the state and transition equations I used were similar to the aforementioned papers, my modeling assumptions were different due to the availability of data.

I used the following dynamic linear model to estimate my latent risk premia of interest:

*Measurement Equation*

\[
y_t = Z\alpha_t + \varepsilon_t, \quad \varepsilon_t \sim N(0, G_t)
\]

Where \( Z = \begin{bmatrix} 1 & \beta_2 & \beta_3 & 1 \\ 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \end{bmatrix} \)

\( \alpha_t = \begin{bmatrix} \text{Expected future average short-rate} \\ \text{Default risk premium} \\ \text{Liquidity risk premium} \\ \text{Segmentation premium} \end{bmatrix} \)

\( y_t = \begin{bmatrix} 10Y \text{ Sovereign Yield} \\ 10Y \text{ EONIA Rate} \\ 10Y \text{ CDS Spread} \\ 10Y \text{ KfW - Bund Spread} \end{bmatrix} \)

\( G_t = \begin{bmatrix} \gamma_1 & 0 & 0 & 0 \\ 0 & \gamma_2 & 0 & 0 \\ 0 & 0 & \gamma_3 & 0 \\ 0 & 0 & 0 & \gamma_4 \end{bmatrix} \)
**Transition Equation**

\[
\mathbf{a}_{t+1} = T \mathbf{a}_t + \mathbf{\eta}_t, \quad \mathbf{\eta}_t \sim N(0, \mathbf{H}_t)
\]

\[
T = \mathbf{I}_4 = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix}
\]

\[
\mathbf{H}_t = \begin{bmatrix} \delta_1 & 0 & 0 & 0 \\ 0 & \delta_2 & 0 & 0 \\ 0 & 0 & \delta_3 & 0 \\ 0 & 0 & 0 & \delta_4 \end{bmatrix}
\]

The measurement equation estimates the value of the matrix of observed variables, \( \mathbf{y}_t \), from the vector of latent risk premia, \( \mathbf{a}_t \), and the transition equation estimates the change in value of the matrix of risk premia \( \mathbf{a}_t \) from time \( t \) to \( t + 1 \).

Following Corradin et al., I assume the following: (a) the error terms for both equations (\( \mathbf{\epsilon}_t \) and \( \mathbf{\eta}_t \)) are normally distributed, (b) there is no intercept for either of the equations, and (c) the transition matrix, \( T \), is set to the identity matrix (since the latent variables are not changing states amongst themselves).

However, while Corradin et al. collect the coefficients (\( \mathbf{\beta} \)) and variance parameters (\( \mathbf{\gamma} \) and \( \mathbf{\delta} \)) together to be estimated using maximum likelihood methods across all countries in their set (i.e. these parameters will be the same for all countries), I opted to estimate unique coefficients and standard deviations for each sovereign by regressing a given country’s sovereign yield on the ten-year EONIA rate, its ten-year CDS spread, and the ten-year KfW-Bund spread.\(^{14}\) Following both Krishnamurthy et al. and Corradin et al., I leave the coefficient on the EONIA rate in the first row of the \( \mathbf{Z} \) matrix equal to 1. This is due to the fact that if we are assuming there is no

\(^{14}\) Note that all coefficients and intercept terms from the OLS regressions of each country turned out to be statistically significant at the 0.05 level.
measurement error, the short-rate and term premium component of the sovereign yield is equal to the EONIA rate. However, in the first row, I set $\beta_2$, the coefficient on the default premium, equal to the coefficient on the CDS spread from the OLS regression of a given sovereign, and $\beta_3$, the coefficient on the liquidity premium, equal to the coefficient on the KfW-Bund spread from the OLS regression. Corradin et al. ended up setting their default premium coefficient equal to 1, finding that this was roughly consistent with their maximum likelihood estimate. However, in my main estimation methodology (I also subsequently discuss an alternative method I use for the $Z$ matrix), I decide to use the coefficients from my regressions instead of the Corradin estimates seeing as (a) my dataset differs from theirs and (b) the OLS coefficients allow for nuances in how much certain premia factor into one country’s sovereign yield as compared to another’s.

Beyond my selection of $\beta_2$ and $\beta_3$, the rest of my $Z$ matrix format mirrors that of the Corradin paper. For instance, I do not estimate a coefficient for the residual segmentation premium but instead assume it to be equal to 1.

**Variance Parameter Selection**

For the matrix of variance parameters for the transition equation, $H_t$, I use the square of the standard errors for each coefficient from a given sovereign’s linear regression, making the aforementioned assumption that the parameters of the regression (i.e. EONIA rate, CDS spread, and KfW-Bund spread) serve as a proxy for the unobserved, latent premia. $[\delta_1, \delta_2, \delta_3, \delta_4]$ therefore correspond to the squared standard errors of the EONIA rate coefficient, the CDS spread coefficient, the KfW-Bund spread coefficient, and the intercept, respectively. I do not use any observed data to directly estimate the segmentation premium (since this premium is considered the residual premium), but the Corradin et al. paper mentions that estimating the $\beta$s

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15 Note that for my data, the set of subscripts [1, 2, 3, 4] correspond to [Short-rate and term premium, default premium, liquidity premium, segmentation premium].
using OLS would lead the intercept term to correspond to the segmentation premium, so I therefore use the intercept’s error in my OLS regression as a proxy for the segmentation premium’s error.

For the matrix of variance parameters for the measurement equation, \( G_t \), I set \( \gamma_1 \) equal to the variance of the residuals of a given country’s OLS regression, seeing as these can serve as a proxy for capturing the transition error in the sovereign yield data. I set \( \gamma_2 = 0 \) since we want to the EONIA rate to exactly reflect the short-rate and term premium and will therefore assume there is no measurement error for this specific latent risk premia. For \( \gamma_3 \) and \( \gamma_4 \), I use the standard deviations of the CDS spreads (\( \gamma_3 \)) and KfW-Bund spreads (\( \gamma_4 \)), continuing under the assumption that these are proxies for the latent premia.

Moreover, another methodological difference from Corradin et al. when implementing the Kalman filter was the choice to make all filter inputs time invariant (Corradin et al. make the measurement error variance matrix time-varying, though they only do this for the CDS spreads and liquidity measures). I find that this is an acceptable methodological deviation due to the fact that the authors themselves note that “the empirical results (…) are not particularly sensitive to adopting an entirely time-invariant measurement error variance matrix” (13).

**Inputs for Initial Estimation of Latent Premia**

Finally, the FKF package in R requires an initial estimation of the latent risk premia, \( a_0 \), as well as the variance of the initial estimations, \( P_0 \). While Corradin et al. specify that they use a non-informative prior distribution, I opt to use initial values corresponding to \( Za_1 \) where I can. For example, as has been specified before, the EONIA rate is a proxy for \( a_1 \), so I use the first observation of the EONIA rate as the value of \( a_{1,1} \). Similarly, the CDS spread and the KfW-Bund spread, scaled by two respective \( \beta \) values, are proxies for the default premium and liquidity.
premium, respectively. I therefore set $a_{2,1}$ equal to the CDS spread measurement at $t = 1$ multiplied by $\beta_2$ from the OLS regression estimate and $a_{3,1}$ equal to the KfW-Bund spread measurement at $t = 1$ multiplied by $\beta_3$ from the OLS regression estimate. Finally, seeing as I do not use any observed data to estimate the residual premium, I set $a_{4,1} = 0$ as a non-informed initial value. For $P_0$, I use the same values of the $\delta$s that I use in the $H_t$ matrix.

**Alternative Estimation of Z Matrix**

Upon initially running the Kalman filter using the $Z$ matrix specified above, my estimates for all eight countries’ liquidity premia turned out to be the same, notably due to the fact that I assumed each one’s liquidity premium to be equal to the KfW-Bund spread (as would be reflected by the coefficient in $Z_{4,3}$ above). For this reason, I also opted to rerun the filter using a different $Z$ matrix (see Appendix B) in order to obtain unique liquidity premia estimates for the eight countries.

There are several differences between this new $Z$ matrix given in Appendix B and the initial one. First, I decide to set the coefficient on the default premium ($\beta_2$ in the initial matrix) equal to 1, implying that the default premium is approximately equal to the CDS spread. I find this to be an acceptable choice, seeing as I use country-specific CDS spread data, which should already account for country-specific variances when being used to measure the default premium.

The main difference, here, is the coefficient on the liquidity premium, which I set to be the reciprocal of $\beta_3$. I do this under the assumption that the liquidity premium is equal to the KfW-Bund spread scaled by some country-specific factor, where in this case the coefficient on the KfW-Bund spread in a given country’s OLS regression ($\beta_3$) is the scaling factor. Reversing the multiplication, the KfW-Bund spread in the $y_t$ matrix would therefore be equal to the
liquidity premium in $a_t$ divided by the scaling factor $\beta_3$ (which is the same as multiplying by $1/\beta_3$ as given in the alternative $Z$ matrix).

I make this choice to attempt to capture a difference between Corradin et al.’s data and mine. To estimate the liquidity premium, Corradin et al. use data from Tradeweb that scales the KfW-Bund spread by a country-specific factor. Seeing as I did not have access to this Tradeweb data and instead used the general KfW-Bund spread, I believe that the aforementioned OLS regression coefficient use allows for a coefficient estimate that adjusts to the given sovereign in lieu of the Tradeweb scaling.

RESULTS

Table 2 reflects the mean risk premia by country across the timeline of observations from January 2014 – December 2021, with the standard deviations reported in parentheses underneath. I report the mean risk premia using both variations of the $Z$ matrix, where $Z_1$ denotes the initial matrix and $Z_2$ denotes the matrix specified in Appendix B. Note all risk premia values are in percentage points.

Figure 1 provides a visual representation of the mean contribution of each risk premia in each sovereign’s yield. This was calculated by taking the absolute value of all the premia on a given day of observation, then dividing the absolute value of a given premium by the sum of all premia on that day. I then took the average of the percent contribution across all days observed to obtain the values in Figure 1.
### Table 2: Mean Risk Premia by Country\(^\text{16}\)

<table>
<thead>
<tr>
<th>Country</th>
<th>Short Rate and Term</th>
<th>Default</th>
<th>Liquidity</th>
<th>Segmentation</th>
</tr>
</thead>
<tbody>
<tr>
<td>All Countries</td>
<td>0.399</td>
<td>Z(_1)</td>
<td>Z(_2)</td>
<td>Z(_1)</td>
</tr>
<tr>
<td></td>
<td>(0.481)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Austria</td>
<td>0.407</td>
<td>0.405</td>
<td>0.237</td>
<td>0.050</td>
</tr>
<tr>
<td></td>
<td>(0.163)</td>
<td>(0.161)</td>
<td>(0.080)</td>
<td>(0.017)</td>
</tr>
<tr>
<td>Belgium</td>
<td>0.562</td>
<td>0.557</td>
<td>0.237</td>
<td>0.131</td>
</tr>
<tr>
<td></td>
<td>(0.246)</td>
<td>(0.237)</td>
<td>(0.080)</td>
<td>(0.044)</td>
</tr>
<tr>
<td>Finland</td>
<td>0.390</td>
<td>0.389</td>
<td>0.238</td>
<td>0.129</td>
</tr>
<tr>
<td></td>
<td>(0.120)</td>
<td>(0.121)</td>
<td>(0.080)</td>
<td>(0.043)</td>
</tr>
<tr>
<td>Italy</td>
<td>1.932</td>
<td>1.929</td>
<td>0.237</td>
<td>0.414</td>
</tr>
<tr>
<td></td>
<td>(0.453)</td>
<td>(0.438)</td>
<td>(0.079)</td>
<td>(0.139)</td>
</tr>
<tr>
<td>Netherlands</td>
<td>0.380</td>
<td>0.378</td>
<td>0.238</td>
<td>-0.035</td>
</tr>
<tr>
<td></td>
<td>(0.139)</td>
<td>(0.139)</td>
<td>(0.080)</td>
<td>(0.011)</td>
</tr>
<tr>
<td>Portugal</td>
<td>1.912</td>
<td>1.901</td>
<td>0.238</td>
<td>0.103</td>
</tr>
<tr>
<td></td>
<td>(1.066)</td>
<td>(1.043)</td>
<td>(0.080)</td>
<td>(0.035)</td>
</tr>
<tr>
<td>Slovenia</td>
<td>1.435</td>
<td>1.440</td>
<td>0.238</td>
<td>0.049</td>
</tr>
<tr>
<td></td>
<td>(0.300)</td>
<td>(0.328)</td>
<td>(0.080)</td>
<td>(0.017)</td>
</tr>
<tr>
<td>Spain</td>
<td>1.145</td>
<td>1.138</td>
<td>0.238</td>
<td>-0.046</td>
</tr>
<tr>
<td></td>
<td>(0.359)</td>
<td>(0.334)</td>
<td>(0.080)</td>
<td>(0.015)</td>
</tr>
</tbody>
</table>

\(^{16}\)All figures are reported in percentage points. Standard deviations are in parentheses.
Figure 1: Mean Contribution of Each Risk Premium by Country

For more output results, please see Appendices C and D. These show time series of the sovereign yields and risk premia results by country using $Z_1$ and $Z_2$, respectively.
DISCUSSION

Sovereign Yield Risk Premia Breakdowns

*Short-Rate and Term Premium*

The short-rate and term premium, estimated to be constant across all countries, accounts for the biggest proportion of the Netherlands, Austria, and Finland’s sovereign yields, at 28.82%, 28.73%, and 27.39%, respectively.

These aforementioned three countries are those with the lowest average sovereign yields and CDS spreads (see Table 1). Seeing as the sovereign yield and CDS spread measures are the only two country-specific measures I input, it makes sense that with low values of their sovereign-specific inputs, the constant short-rate and term premium takes up a greater share of the overall yield.

This finding is also consistent with comparable countries in the Corradin et al. paper. Corradin et al. look at France and Germany, which are similar nations in the sense of lower yields and high credit ratings. The short-rate and term premium for France and Germany accounts for a large portion of these nations’ yields\(^{17}\), comparatively more than Spain and Italy (the more “risky” countries in the paper’s sample).

*Default Premium*

Portugal, Italy, and Spain have the highest average default risk premia, at 57.44%, 53.25%, and 51.56%, respectively. This makes sense in light of how these three countries are part of the GIIPS\(^{18}\) group that faced significant hardships during the European Debt Crisis. Moreover, these three sovereigns have the lowest credit ratings relative to the other five, with

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\(^{17}\) 45% and 46%, respectively (Corradin et al., 28).

\(^{18}\) Greece, Ireland, Italy, Portugal, and Spain
Spain’s S&P rating being A and Portugal and Italy’s being BBB.\textsuperscript{19} These ratings imply that investors should require additional return from these sovereigns’ bonds to compensate for their comparatively higher default risk.

On the other hand, the countries with the three lowest default risk premia are Finland, the Netherlands, and Austria, at 23.93\%, 24.28\%, and 25.84\%, respectively. This is consistent with the fact that these three countries have the highest credit ratings, with their S&P ratings being AAA (the Netherlands) and AA+ (Finland and Austria).\textsuperscript{20}

\textbf{Liquidity Premium}

As found in the previous studies, the liquidity premium is consistently the lowest proportion of any given sovereign’s yield. The sovereign for which this premium is the highest is Italy, at an average of 11.67\%. The liquidity premium seems to take up the smallest fraction of Slovenia’s yield, at an average of 1.674\%.

For the $Z_2$ matrix results, the liquidity premia for both Spain and the Netherlands turn out to be consistently negative. Unfortunately, there does not seem to be any explanation in the literature regarding how or why this could be the case. This likely error in estimation may therefore be due to the methodology, the data, or both.

Evidence that highlights the spuriousness of these results includes the recent finding that during the COVID-19 pandemic, "European sovereign bond markets experienced a strong deterioration of market liquidity (...) [and that] these patterns were similar across the largest three euro area economies" (Moench, Pelizzon, Schneider 2021). Given that the largest economies faced liquidity crunches, one would assume a negative liquidity premium for all countries, and the smaller ones even more so. The $Z_1$ matrix results do support the Moench et al. findings,

\textsuperscript{19} European Parliament.
\textsuperscript{20} European Parliament.
seeing as the liquidity premium spikes and reaches a maximum in early 2020 (see Appendix C). However, in the $Z_2$ results, both Spain and the Netherlands exhibit a further decline in their already negative liquidity premia in the early 2020s (see Appendices D5 and D8), evidencing a high probability of there being some sort of issue in the results.

**Segmentation Premium**

The segmentation premium is greatest for the Netherlands, Austria, and Finland (44.49%, 41.85%, and 40.15%, respectively) and smallest for Italy, Spain, and Portugal (25.07%, 22.66%, and 21.43%, respectively). Belgium and Slovenia fall in the middle of these two sets. As with the other risk premia, there seems to be a recurring trend in which countries exhibit similar breakdowns of risk premia.

Moreover, similar to the Corradin et al. paper, the segmentation premium accounts for a greater proportion of the more stable countries’ yields relative to the less stable countries’ yields. Compared to Corradin et al.’s estimates of 11% and 19% for the segmentation premia of Italy and Spain, respectively, my estimates are higher. However, it is important to note that seeing as the segmentation premium is estimated as a residual, it could also account for other risk premia not directly accounted for. It is therefore hard to infer with any confidence whether the segmentation premium estimated is truly a reflection of segmentation risk and whether its magnitude is accurate.

Another interesting result from the segmentation premium\(^{21}\) is that it is the only latent variable that fluctuates between negative and positive values. Belgium and Italy’s segmentation premia are positive during 2014, and Portugal and Spain’s are as well but with the addition of being briefly positive again in later years (see Appendices C2, C4, C6, and C8). According to

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\(^{21}\) These results appear when using the $Z_1$ matrix (i.e. with a constant liquidity premium for all countries). When using $Z_2$, only Spain’s segmentation premium is briefly positive during 2014.
Corradin et al., “a consistently negative segmentation premium means that investors are willing to accept a lower return from sovereign bonds compared to holding an alternative position that has the same (or similar) payoffs” (24). Perhaps the positive segmentation premia for certain countries in 2014, which imply that investors received negative non-pecuniary value from holding these bonds, are due to the lingering effects of the European Debt Crisis. However, as mentioned before, due to the residual nature of this premium, it is difficult to infer whether this was due to segmentation risk or due to perhaps the redenomination risk being high at this point in time. Neither Krishnamurthy et al. or Corradin et al.’s analysis includes 2014, so there is no available information in the previous literature regarding the relative values of the segmentation and redenomination premia during this year.

Limitations of Risk Premia Selection

The decomposition of a sovereign yield into different premia is an inherently subjective endeavor, seeing as it depends on the researcher’s beliefs of what premias are relevant. For example, while Krishnamurthy et al. did not estimate a liquidity premium, Corradin et al. did.

I did not estimate a redenomination premium, while both previous papers did and found it to be a sizeable portion of countries’ sovereign yield breakdowns. For this reason, one of the main issues with my segmentation premium estimate may be that it also accounts for the redenomination risk premium. Moreover, seeing as I used a linear regression to estimate the coefficients and relative weights of the various risk premia in the sovereign bond yields, adding another premium would likely affect the weights of all the other premia and change my results as a whole.
Limitations of Data

One limitation of my study was the inability to access certain financial data that could help inform country-specific estimates of risk premia. For instance, the data for the ISDA basis is difficult to find and only available for a very limited number of countries on Bloomberg. For the “quanto CDS” method of measuring the redenomination premium, I was unable to find any EUR CDS quotes on Bloomberg. I also did not have access to Tradeweb and consequently could not observe country-specific liquidity premia measures for my Kalman filter inputs.

Moreover, the KfW-Bund data generally proved to be problematic, notably since the liquidity risk premia for the Netherlands and Spain turned out negative using $Z_2$. These doubtful results likely stem from how the KfW-Bund spread is an imperfect measure of liquidity risk. First, liquidity relates more closely to the volume of bonds traded and the frequency of trading activity, which the KfW-Bund spread fails to capture. The operationalization of liquidity through the KfW-Bund spread also relies on the assumption that the only premium that would result between a KfW bond and German Bund would be due to liquidity risk. However, there could be other unobserved premia that explain differences between the two.

A more ideal way to measure the liquidity premium would be to look at the daily volume of sovereign bonds traded by country, then to find the percent composition of one sovereign’s bond trade volume relative to the total volume of sovereign bonds traded on a given day. This percent could then be used as a scaling factor for some conventional liquidity premium measure, such as the KfW-Bund spread or Bund bid-ask spread. However, this sort of data on volume traded is quite difficult to find, if even aggregated or publicly available at all.
Limitations of Kalman Filter

Z Matrix

When deciding on values for the $Z$ matrix, there was partial reliance on assumptions from Corradin et al. and Krishnamurthy et al. (e.g. the EONIA rate being equal the short-rate and term premia) and partial reliance on values from the OLS. This inconsistency of dependence on different methods could therefore lead to a mathematically incorrect implementation of the filter on the data. For example, for element (1, 2) of the $Z_2$ matrix (the estimate of the coefficient on the default premium), I used Corradin et al.’s estimate of 1, yet for (1, 3), the coefficient on the liquidity premium, I used the estimate from my OLS regression.

An additional source of inconsistency was between $Z_1$ and $Z_2$, where in the first row of $Z_1$, I put $\beta_3$ (i.e. an estimated coefficient for the liquidity premium) as the Corradin paper did. In $Z_2$, however, I put the reciprocal, $1/\beta_3$, instead. I opted to use the reciprocal $1/\beta_3$ as the scaling factor (as opposed to $\beta_3$), seeing as in the OLS, $\beta_3$ multiplied by the KfW-Bund spread is one of the terms in the summation of the sovereign yield, so reversing the multiplication, the liquidity premium multiplied by the reciprocal of $\beta_3$ should equal the KfW-Bund spread. In hindsight, I should have been more consistent in either using the coefficient or the reciprocal in my calculations.

The assumptions and choices made in coding the $Z$ matrix were likely the biggest weakness in the implementation of the Kalman filter. A potential solution to this issue would have been to attempt to directly replicate Corradin et al.’s matrix by estimating the coefficient

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22 At position (1, 3).
23 Sovereign Yield = $\beta_1$(EONIA Rate) + $\beta_2$(CDS Spread) + $\beta_3$(KfW-Bund Spread) + Intercept, which is used to estimate: Sovereign Yield = Short-Rate and Term Premium + Default Premium + Liquidity Premium + Segmentation/Residual Premium
parameters through maximum likelihood, then using the same loading coefficients that they did. However, seeing as this paper a) estimated the coefficients using OLS and b) used a country invariant estimate of liquidity, deviating from their matrix methodology could be considered an appropriate discretionary choice.

**Coefficient Estimates**

The $\beta$ coefficients values in the Kalman filter are important, seeing as they help calculate the observed data from the latent risk premia. The parameters used as inputs into the model could not have been the best ones, notably due to the fact that they were estimated using a Classical OLS regression.

One potential slight improvement on the Classical OLS would have been to do a Bayesian version instead. This would allow for posterior sampling of $\beta$ values to see which coefficients could be better fits to the data. However, the biggest issue in implementing OLS here, regardless of whether Classical or Bayesian, is that some of the fundamental assumptions of OLS do not hold for the data used in this paper. For instance, there is definitely an issue of multicollinearity between the three parameters$^{24}$ and the sovereign yield, seeing as most macroeconomic financial variables related to a given sovereign are likely to have some sort of relationship with each other. Moreover, there are definitely issues of autocorrelation. Running Durbin-Watson tests on the eight fitted OLS models by country resulted in statistically significant evidence of autocorrelation (see Appendix E). This makes sense seeing as all the data were time series.

For this reason, while I opted to use OLS estimations of the $\beta$ values as a simplification, again, perhaps it would have been best to follow Corradin et al.’s method of maximizing the

$^{24}$ EONIA rate, CDS spread, and KfW-Bund spread.
likelihood of the parameters across all the data. The one disadvantage of Corradin et al.’s method, however, is that it does not allow for country-specific deviations in coefficients, which are likely to exist.

**Areas of Further Study**

**Larger Sample Size**

It would be ideal to analyze all Euro sovereigns using the Kalman filtering methodology, as well as to estimate all risk premia covered in the literature. However, this would be contingent on the availability of data. It remains to be seen whether future papers will find ways to operationalize some of the variables that have been difficult to estimate, notably the liquidity and redenomination risk premia.

**Event Study**

After extracting the latent risk premia, Corradin et al. also performed an event study to see how the different risk premia and sovereign yields reacted to certain fiscal and monetary policy announcements. This event study application is therefore an area of potential further extension, where one could look at how yields and risk premia move in response to different key events. These events could not only include economic ones, but also social and political ones (e.g. national elections).

**CONCLUSION**

From this analysis, I find that the default premium and segmentation premium are the two main components of sovereign yields, while the liquidity premium is consistently the smallest component. The segmentation premium tends to be a greater share of more stable countries’ sovereign yields, while the default premium tends to be a greater share of less stable countries’
yields (i.e. the GIIPS countries in my sample). However, seeing as the segmentation premium was estimated as a residual, it follows that it may also capture other risk premia unaccounted for, namely the redenomination risk premium.

The countries in my sample exhibit similar patterns by group. Countries with lower average sovereign yields and higher credit ratings tend to fall into one group (the Netherlands, Finland, and Austria), while countries with higher average sovereign yields and lower credit ratings fall into another (Spain, Portugal, and Italy). Belgium and Slovenia would comprise a middle group, showing results in between the others.

The main challenge with the implementation of the Kalman filter turned out to be the estimation of the coefficients for the loading matrix, $Z$. The output results were highly sensitive to the inputs for $Z$. While easy to implement, the use of OLS regressions to estimate the $\beta$ parameters for the $Z$ matrix had limitations in terms of reliability.

Going back to the initial question that got me interested in this research topic, what can these results tell us about the economic and financial convergence of the Euro area countries? The trends uncovered in different groups of countries imply that the convergence of yields and their premia has yet to occur, if ever. Countries differ in which risks their debt investors perceive to exist, emphasizing the need for the ECB’s monetary policy, as well as nation-specific fiscal policy, to be able to address a broad range of financial concerns, ranging from issues of liquidity to issues of sovereign trust and creditworthiness. However, looking at the yields and their premia since early 2020 when the COVID-19 pandemic started, it seems that the supranational Euro-area response to the imminent threat of another big financial crisis was fairly effective, seeing as sovereign yields at the end of 2021 are lower than their early 2020 peak, especially the default risk premium component. It remains to be seen whether this downward trend will continue, and
which premia will bear what weights, as the economic and financial landscape continues to change.
APPENDIX A

Bloomberg Data Tickers

<table>
<thead>
<tr>
<th>Country</th>
<th>EONIA Rate</th>
<th>KiW-Bund Spread</th>
<th>Sovereign Yield</th>
<th>CDS Spread</th>
</tr>
</thead>
<tbody>
<tr>
<td>All Countries</td>
<td>EUSWE10 Currency</td>
<td>IB10KFWBVLI</td>
<td>GT [COUNTRY CODE] 10Y GOVT</td>
<td>[COUNTRY CODE] CDS USD SR 10Y D14 Corp</td>
</tr>
<tr>
<td>Austria</td>
<td>RAGB 0.9 02/20/2032</td>
<td>AUST CDS USD SR 10Y D14 Corp</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Belgium</td>
<td>BGB 0.35 06/22/32</td>
<td>BELG CDS USD SR 10Y D14 Corp</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Finland</td>
<td>RFGB 0 1/8 09/15/31</td>
<td>FINL CDS USD SR 10Y D14 Corp</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Italy</td>
<td>BTPS 0.95 06/01/32</td>
<td>ITALY CDS USD SR 10Y D14 Corp</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Netherlands</td>
<td>NETHER 0 ½ 07/15/32</td>
<td>NETHER CDS USD SR 10Y D14 Corp</td>
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<td>Portugal</td>
<td>PGB 1.65 07/16/2032</td>
<td>PORTUG CDS USD SR 10Y D14 Corp</td>
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</table>

APPENDIX B

Alternative Z Matrix for Kalman Filtering

\[
Z = \begin{bmatrix}
1 & 1 & \frac{1}{\beta_3} & 1 \\
1 & 0 & 0 & 0 \\
0 & 1 & 0 & 0 \\
0 & 0 & \frac{1}{\beta_3} & 0 
\end{bmatrix}
\]
APPENDIX C

Sovereign Yield Decomposition Using Z₁

Figure C1: Austria

Figure C2: Belgium

Figure C3: Finland

Figure C4: Italy
APPENDIX D

Sovereign Yield Decomposition Using $Z_2$

Figure D1: Austria

Figure D2: Belgium

Figure D3: Finland

Figure D4: Italy
Figure D5: Netherlands

Figure D6: Portugal

Figure D7: Slovenia

Figure D8: Spain
APPENDIX E

Durbin-Watson Test Results for OLS Models

<table>
<thead>
<tr>
<th>Country</th>
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<td>Belgium</td>
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<tr>
<td>Finland</td>
<td>0.029</td>
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<tr>
<td>Italy</td>
<td>0.084</td>
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<tr>
<td>Netherlands</td>
<td>0.042</td>
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<tr>
<td>Portugal</td>
<td>0.082</td>
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<td>Slovenia</td>
<td>0.060</td>
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<tr>
<td>Spain</td>
<td>0.119</td>
</tr>
</tbody>
</table>

All results are statistically significant at the 0.05 level.
REFERENCES


