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Abstract

This dissertation is organized in four chapters. Chapter 1 introduces the methodology used and briefly describes each chapter. Chapters 2 and 3 test for the existence of fractional cointegration relationships with applications to macroeconomic and financial time series. Chapter 4 studies cross-country monetary spillover effects within the Euro Area.

Chapter 2 revisits the question whether volatilities of different markets and trading zones have a long-run equilibrium in the sense that they are fractionally cointegrated. The notion of fractional cointegration allows for long-term equilibria with a higher degree of persistence than allowed for in the standard cointegration framework. We consider the U.S., Japanese and German stock, bond and foreign exchange market to see whether there is fractional cointegration between the markets in one trading zone or for one market across trading zones. Also, the other combinations of different markets in different trading zones are considered. Applying a purely semiparametric approach through the whole analysis shows that fractional cointegration can only be found for a small minority of different cases. Investigating further we find that all volatility series show persistence breaks during the observation period which may be a reason for different findings in previous studies.

Chapter 3 tests for convergence of EMU inflation rates and industrial production growth rates by using the same semiparametric fractional cointegration methodology as in Chapter 2. Over the full sample from 1999:01-2019:12, we find evidence of fractional cointegration in both inflation and industrial production among many country pairs, where nominal convergence does not necessarily imply real convergence and *vice versa*. Our results suggest some evidence for "convergence clusters" among either core or periphery countries in the case of inflation. By contrast, we find more evidence of mixed core-country cointegration pairs for industrial production, where convergence may be driven by trade links. Testing for a break in the persistence structure, the results show evidence of a break in the persistence of both inflation and industrial production during the beginning of the financial crisis in a number of countries. In all cases, persistence is substantially higher after the break, suggesting a higher potential for diverging processes during economic crises. However, in some cases we find a second break marking the end of the crisis period, with persistence back at pre-crisis levels after the second break.

Motivated by the results in Chapter 3, Chapter 4 extends the Eurozone framework by constructing a large dataset and investigating the effects of monetary policy on the economy of 12 country members of the Euro Area. The dynamic relations among monetary policy, real economy, prices and the financial market are examined using time-varying parameter factor-augmented VAR models for the period 2000:01-2019:12. Although the transmission mechanism has demonstrated success in the eurozone, we still find asymmetries over time and across core and periphery countries in both real and nominal terms.

Keywords: European monetary policy, European financial crisis, Macroeconomic convergence, High-frequency data, Realized volatility, Semiparametric Estimation, Fractional Cointegration, Persistence breaks, TVP-FAVAR

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Chapter 1

Introduction

Time series are used extensively in empirical research in both areas of macroeconomics and finance. One of the main applications of time series analysis is to capture the "co-movement" between the variables of interest over time. For that purpose and due to the rapid development in statistical analysis, economists and policy makers are constantly using time series modelling to capture such relationships which contribute to monetary and fiscal policy designing, business planning, financial asset and risk management, among others.

This dissertation presents three essays on applied time series by using advanced techniques to address recent macroeconomic and financial issues. In the first two essays, an extension of the standard cointegration methodology is used. Cointegration was first introduced by Engle and Granger (1987) as time series that wander extensively and yet there can be forces which tend to keep them not too far apart. Thus, cointegration is considered as a long-run equilibrium relationship which involves time series with integer orders of integration. Fractional cointegration is a more relaxed definition than standard cointegration's which allows the order of integration to be a fraction. Formally, if two processes are integrated of the same fractional order and there exists a linear combination of them with a lower order of (fractional) integration then the series are said to be fractionally cointegrated.

The flexibility of the noninteger integration has led to the development of several testing and estimation procedures for the existence of a fractional equilibrium relationship. Focusing on the semiparametric approaches where no short-run dynamic specifications are required, some of the most popular procedures in the literature are those of Chen et al. (2006), Hualde and Velasco (2008), Marmol and Velasco (2004), Nielsen and Shimotsu (2007), Nielsen (2010), Robinson and Yajima (2002), Robinson (2008), Souza et al. (2018), Wang et al. (2015), Zhang et al. (2019). However, some of the procedures are restrictive enough and their selection depends on the empirical application.

Chapter 2 examines whether volatilities of different financial markets and trading zones have a long-run equilibrium by using a variety of the semiparametic tests indicated above. This essay is motivated by a previous work of Clements et al. (2016) where they use high-frequency data (5 minutes) for foreign exchange rates, stocks and bonds for the period 1999-2015 and find evidence of common trends that drive volatility in global markets. First, we estimate the order of integration of the series under examination and we find either low or not equal order of integration between volatilities. This is in line with other findings in the literature. Wenger et al. (2017) find lower orders of integration which makes the concept of fractional cointegration questionable for these asset classes. Nguyen et al. (2020) find different orders of integration for different markets. Therefore, the key assumption of fractional cointegration for equal orders of integration among the volatility series cannot be fulfilled and is excluded for many cases.

Next, we use the semiparametric testing procedures by Chen et al. (2006), Souza et al. (2018) and Wang et al. (2015) and estimate the order of integration of the linear combination for the case-pairs we are allowed to. Only a few cases reject the null hypothesis of no fractional cointegration. Hence, we cannot support the findings by Clements et al. (2016). Finally, we extend our analysis by testing for potential breaks in the order of integration. Recent studies focus on the impact of financial crises on the volatility spillover. We therefore test for the existence of structural breaks and find evidence of breakpoints in all series during the global financial crisis and European debt crisis. We, then, apply rolling window regressions to gain further information of the persistence dynamics, which indicates a change of persistence over time. Therefore, we re-apply the persistence testing according to the estimated breakpoints and observe shifts in the order of integration during different periods. Specifically, we find lower volatility persistence after the breaks. This can result in spurious results on fractional cointegration if neglected in the analysis. We therefore commit to taking into account the frequent changes in persistence over time.

Chapter 3 uses the same methodology as in Chapter 2 focusing now on the macroeconomic convergence of the Eurozone. The recent European debt crisis following the crisis of 2008 revealed potential macroeconomic divergence within a monetary union (Borio and Disyatat (2011),Gnimassoun and Mignon (2016)). In this essay, we thus use fractional cointegration methods as a tool to measure nominal and real convergence or divergence since the start of European Monetary Union. Therefore, we are using monthly observations on inflation growth rates and quarterly observations on industrial production rates for the cases of Austria, Belgium, Finland, France, Germany and Netherlands ("core countries"), as well as for Greece, Ireland, Italy, Portugal and Spain ("periphery countries"). Our analysis starts at 1999:01, beginning of the European bank acting as a "single" central bank until 2019:12.

For the case of inflation rates, we find evidence of fractional integration in all countries, as well as fractional cointegration relationships among the country pairs. Specifically, we find stronger evidence among core and periphery countries rather than in mixed core-periphery countries. This could lead to current account disequilibria if the development is sustained over longer periods of time. For the case of industrial production rates, we also find evidence of fractional integration and cointegration. There is evidence in core country-pairs, mixed core-periphery pairs and fewer long-run equilibria among periphery countries. This could be an indication of real convergence processes being driven by trade links among the EMU countries.

As a next step, persistence break tests in the order of integration of all series are conducted. Both inflation and industrial production rates present breakpoints during the financial crisis for the majority of the countries. A second breakpoint is found for a few cases during 2011-2015 marking the end of the crisis. Overall, our results show that real convergence does not necessarily imply nominal convergence. Despite the financial and

European debt crisis, we still have evidence that the crisis period ended before the end of our sample in some cases and fractional cointegration relationships still exist for many pairs for the full sample period.

Through further research on co-movement of economic phenomena over time, Chapter 4 is using a time-varying parameter Factor Augmented Vector Autoregressive (FAVAR) model to study the spillover effects of monetary policy on Euro Area countries. FAVAR models were developed by Bernanke et al. (2005) as a method that summarizes the dynamics of a large set of information to a small number of factors in the standard Vector Autoregressive (VAR) model. The small set of the estimated factors along with an observed policy variable reduce the ommited variable bias problem and the information insuffiency that is likely to be reflected in the standard VAR approach. Due to the significant information exploit the model offers, it is widely used in the identification of the monetary transmission mechanism.

The analysis on Chapter 4 builds on a recent strand of literature that observes asymmetric spillover on Euro Area countries since the beginning of EMU (Barigozzi et al. (2014), Blaes (2009),Boivin et al. (2008), Galariotis et al. (2018), Laine (2020), Peersman (2004), Potjagailo (2017)) and mainly during and after the financial crisis where nominal interest rates declined globally. By employing a large dataset of macroeconomic and financial series in a FAVAR model with time varying coefficients and stochastic volatility, we contribute to the understanding of the size and direction of European monetary policy spillovers over time. Impulse response analysis strengthens the evidence of asymmetric transmission for the cases of inflation, unemployment rates and long term government bond rates between core and periphery countries.

The rest of the dissertation is organized as follows. Chapters 2 to 4 describe each of the three essays as outlined above. Chapter 5 lists the bibliographic references.

Chapter 2

Volatility transmission across financial markets: A semiparametric analysis

Co-authored with Mwasi Mboya and Philipp Sibbertsen.

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Chapter 3

Measuring Macroeconomic Convergence and Divergence within Economic and Monetary Union Using Long Memory

Co-authored with Lena Dräger and Philipp Sibbertsen.

Under revision for Studies in Nonlinear Dynamics & Econometrics.

3.1 Introduction

Ever since the idea of a European Monetary Union (EMU) was born, economists have discussed the need for member countries to converge in nominal and real terms in order to ensure the stability of the union. This follows from the concern that inflation differentials between member countries with a single monetary policy may lead to differences in real interest rates and in international competitiveness, thereby creating real divergence as measured in current account imbalances. These imbalances are difficult to address if the affected countries share a common currency and thus have permanently fixed nominal exchange rates and a single monetary authority Mundell (1961).

Early studies before the start of EMU demonstrate some success in terms of nominal convergence of the member states Beliu and Higgins (2004), whereas others use cointegration analysis to demonstrate potential long-run stability problems with respect to macroeconomic dynamics in the so-called "periphery" countries Italy, Spain and Portugal Haug et al. (2000). However, the initial convergence in inflation rates was somewhat reversed after the start of EMU, resulting in persistent inflation differentials, where groups of countries showed inflation rates persistently above or below the EMU average ECB (2003). On the one hand, such differences may imply that countries with below-average per-capita real GDP levels are experiencing a catching-up growth process with temporary above-average real growth and inflation rates Blanchard and Giavazzi (2002). As long as the catching-up process leads to convergence in productivity levels, it may be regarded as a temporary phenomenon. On the other hand, however, persistent differences in inflation rates within a monetary union impair the efficiency of monetary policy. Since the ECB can only target average inflation within the union, in such a situation monetary policy may be overly expansive in regions with above-average inflation rates. Hence, these regions experience real appreciation, while regions with below-average inflation rates experience real depreciation. The loss in international competitiveness of regions within the monetary union and the resulting current account deficits bears the danger of potentially large economic costs in case of a current account adjustment Borio and Disyatat (2011); Gnimassoun and Mignon (2016). Indeed, the economic crises following the financial crisis of 2008 and the European sovereign debt crisis after 2010 led to a real depreciation in the "periphery" countries through deflation with large losses in real GDP and employment.

Since nominal and real convergence is difficult to identify empirically, in this paper we use fractional cointegration methods accounting for long-memory equilibrium processes as a tool to measure nominal and real convergence or divergence since the start of EMU. In this set-up, economic convergence does not necessarily imply convergence in levels. Rather, the existence of a stable long-run equilibrium between macroeconomic variables across EMU countries is regarded as evidence of close and stable macroeconomic relationships among those countries. This is what we term macroeconomic convergence. Our definition of macroeconomic convergence is thus more flexible than convergence in levels. We argue that the flexible framework accounts for potentially persistent deviations from the long-run equilibrium. Importantly, the existence of a long-run equilibrium nevertheless implies that the economies of the currency union with fractional cointegration relationships will eventually return to their stable long-run equilibrium.

Testing for the existence of a long-run equilibrium with fractional cointegration has the advantage of allowing the degree of integration to take any real number on the unit interval. Therefore, the framework allows for higher flexibility by allowing the long-run equilibrium to be a mean-reverting rather than a short-memory stationary process. Evidence for fractional cointegration with long memory in the equilibrium errors then indicates the existence of a long-run equilibrium, while at the same time deviations from this equilibrium may be very persistent. In contrast to the previous literature, we account for the full EMU period including the financial crisis period from 2008 onwards. However, this makes our analysis vulnerable to spurious results through changes in the convergence mechanism and thus in the cointegrating relation. We therefore apply the regression-based Lagrange Multiplier test introduced by Martins and Rodrigues (2014) to test for structural breaks in the order of integration and estimate the break dates.

We estimate the degree of fractional long memory for inflation and industrial production growth rates (IP) in 11 EMU countries for the period from January, 1999, to December, 2019, where we distinguish between "core" and "periphery" countries. Our results suggest breaks in the persistence of both inflation and industrial production around the beginning of the financial crisis in 2007-08 for a majority of countries in our sample. In some countries, we find a second break marking the end of the crisis period. Inflation and industrial production are generally found to be more persistent in the periphery countries than in the core countries. In addition, persistence is estimated to be higher in the crisis period. This implies that both inflation rates and industrial production were more integrated in the EMU before the financial crisis of 2008, while the higher persistence during the crisis period carries the danger of longer-lasting divergence in case of asymmetric shocks.

We further test for the existence of fractional cointegration relationships for inflation and IP among all country pairs of our sample and estimate the memory of the residuals d_v by using Narrow Band Least Squares estimation for the cointegrating vector β. We find evidence of fractional cointegration for both inflation and industrial production over the whole sample period. In the case of inflation, we find evidence of fractional cointegration among core and among periphery countries, but fewer evidence of mixed core-periphery cointegration pairs. This result suggests a certain risk of "convergence clubs" regarding inflation, which could lead to current account disequilibria if the development is sustained over longer periods of time. By contrast, in the case of industrial production, we find evidence of cointegration among core country-pairs, but also among mixed core-periphery pairs, with fewer long-run equilibria among periphery countries. In summary, it seems that real convergence does not necessarily imply nominal convergence and *vice versa*.

Our paper builds on previous work by some of the authors in Leschinski et al. (2021). Using the fractional cointegration methodology, Leschinski et al. (2021) provide empirical evidence for periods of convergence and divergence for long-term EMU government bonds that coincide with bull- and bear-market periods in the stock market. Specifically, stronger market integration is associated with bull-market periods and is more intense among core countries than among periphery countries. Periods of disintegration coincide with bear-market periods. Their results thus imply time-variation in the degree of convergence of EMU government bonds, with the possibility of divergence even before the financial crisis of 2008. Moreover, the authors report evidence of disintegration in government bonds for all countries during the period of the financial crisis and the European sovereign debt crisis from 2008 onwards.

Our paper further relates to the previous literature using fractional cointegration methods to measure macroeconomic convergence between European economies. In an early study, Beliu and Higgins (2004) use fractional cointegration tests to evaluate the convergence of inflation, long-term interest rates and industrial production of 14 EU countries vis-à-vis Germany. The sample (1957-2001) covers the period until the Euro cash changeover, whereas we focus on the period after the start of EMU in 1999. The authors present evidence of nominal convergence in inflation and long-term interest rates, as these series are fractionally cointegrated with the German counterpart. However, the equilibrium errors display long memory, so that any deviation from equilibrium will be persistent. The authors find no evidence of fractional cointegration in industrial production and, thus, no evidence for real convergence. Meller and Nautz (2012) test for differences in inflation dynamics among European countries before and after EMU using panel estimates of fractional cointegration. Their results suggest that inflation persistence converged and was significantly reduced with the introduction of EMU. More recently, Hualde and Iacone (2017b) analyze both level inflation as well as inflation differentials between EMU country pairs allowing for long memory and cointegration with the test procedure derived in Hualde and Iacone (2017a). Their results suggest that the "core" economies of EMU are more integrated than the "periphery" countries, as the latter show more persistent inflation differentials. Similarly, Karanasos et al. (2016) study convergence of inflation among EMU countries for the period 1980-2013 using a broad range of test methods, which includes tests allowing for long memory and for structural breaks. Similar to the study by Hualde and Iacone (2017b), the authors present evidence for three clubs of convergence consisting of "core" EMU countries, while there is evidence of divergence in inflation for the remaining countries. We extend these previous studies, as we account for the effects on nominal and real convergence of the recent period of economic turmoil following the financial crisis in 2008. In addition, we compare the results from several semiparametric tests for fractional cointegration by Souza et al. (2018), Wang et al. (2015) and Chen and Hurvich (2006).

Other studies applying the concept of fractional cointegration evaluate, for instance, the stability of money demand functions Caporale and Gil-Alana (2005), the effect of inflation targeting regimes on inflation persistence Canarella and Miller (2017) or the effect of a monetary policy shock when long memory in the output gap and inflation is accounted for Lovcha and Perez-Laborda (2018).

The remainder of the paper is structured as follows. The econometric methodology of fractional cointegration with long memory is detailed in section 3.2. Section 3.3 describes the data set. The results of our anaylsis are presented and discussed in sections 3.4 and

3.5. Finally, section 3.6 summarizes and concludes.

3.2 Methodology

Inflation and industrial production persistence in Europe has attracted renewed attention since the foundation of EMU and the implementation of a common monetary policy among the member countries. Persistence is here measured by the order of integration of the inflation series. If this order of integration d fulfills $0 < d < 0.5$, the series is said to exhibit stationary long memory or stationary long-range dependence resulting in a hyperbolic decay of the autocorrelation function and therefore in a slow decay of dependencies between observations far away from each other. Also the impulse-responses show a hyperbolic decay. For $0.5 < d < 1$, the series is non-stationary, but still mean reverting in the sense that the expected time for returning to its mean is finite. Therefore, the concept of long memory or fractional integration widens the traditional $I(1)/I(0)$ duality by allowing for highly persistent equilibrium processes.

A significant number of empirical studies that have been adopted in recent years support the existence of long-range dependence in inflation rates. Baillie et al. (2002), Canarella and Miller (2016), Canarella and Miller (2017) Hassler and Wolters (1995), Kumar and Okimoto (2007) and Meller and Nautz (2012) among others, measure persistence as a fractionally integrated process. Moreover, fractional integration techniques are able to capture convergence among inflation policies and detect persistence shifts more consistently than the $I(1)/I(0)$ approach, as assumptions about the $I(0)$ process of inflation process or the short-run persistence structure are not required (Kumar and Okimoto (2007)). Inflation convergence requires the existence of a stable equilibrium relationship, but not exact equality between the inflation rates. A similar definition of convergence can also be applied to real economic variables, such as industrial production (IP).

In general, (fractional) cointegration is an equilibrium concept where the persistence of the cointegrating residual d_v determines the speed of adjustment towards the cointegration equilibrium $\beta' X_t$, and shocks have no permanent influence on the equilibrium as long as $d_v < 1$ holds. We therefore allow for fractional cointegration and consider a bivariate system of the form

$$X_{1t} = c_1 + \xi_1 Y_t + \Delta^{-(d-b_1)} u_{1t} \mathbb{1}_{\{t>0\}} \tag{3.1}$$

$$X_{2t} = c_2 + \xi_2 Y_t + \Delta^{-(d-b_2)} u_{2t} \mathbb{1}_{\{t>0\}} \tag{3.2}$$

$$Y_t = \Delta^{-d} e_t \mathbb{1}_{\{t>0\}}, \tag{3.3}$$

where the coefficients c_1, c_2, ξ_1, and ξ_2 are finite, $0 \leq b_1, b_2 \leq d$, L is the lag-operator, the fractional differences $\Delta^d Y_t = (1-L)^d Y_t$ are defined in terms of generalized binomial coefficients such that

$$(1-L)^d = \sum_{k=0}^{\infty} \binom{d}{k} (-1)^k L^k = \sum_{k=0}^{\infty} \pi_k L^k,$$

$$\text{with} \quad \binom{d}{k} = \frac{d(d-1)(d-2)\dots(d-(k-1))}{k!},$$

and e_t and $u_t = (u_{1t}, u_{2t})'$ are martingale difference sequences. The memory of both X_{1t} and X_{2t} is determined by Y_t so that they are integrated of the same order d, denoted by $X_t \sim I(d)$, where the memory parameter is restricted to $d \in (0, 1]$ and $X_t = (X_{1t}, X_{2t})'$. Since it is assumed that $u_{1t} = u_{2t} = e_t = 0$ for all $t \leq 0$, the processes under consideration are fractionally integrated of type-II. For a detailed discussion of type-I and type-II processes we refer to Marinucci and Robinson (1999). The spectral density of X_t can be approximated by

$$f_X(\lambda) \sim \Lambda_j(d)\, G\, \overline{\Lambda_j(d)}, \quad \text{as } \lambda \to 0^+, \tag{3.4}$$

where G is a real, symmetric, finite, and non-negative definite matrix, $\Lambda_j(d) = \text{diag}\left(\lambda^{-d} e^{i\pi d/2}, \lambda^{-d} e^{i\pi d/2}\right)$ is a 2×2 diagonal matrix and $\overline{\Lambda_j(d)}$ is its complex conjugate transpose. The periodogram of a process X_t is defined through the discrete Fourier transform $w_X(\lambda_j) = \frac{1}{\sqrt{2\pi T}} \sum_{t=1}^{T} X_t e^{i\lambda_j t}$ as $I_X(\lambda_j) = w_X(\lambda_j)\overline{w_X(\lambda_j)}$, with Fourier frequencies $\lambda_j = 2\pi j/T$ for $j = 1, ..., \lfloor T/2 \rfloor$, where the operator $\lfloor \cdot \rfloor$ returns the integer part of its argument.

The two series X_{1t} and X_{2t} are said to be fractionally cointegrated, if there exists a linear combination

$$\beta' X_t = v_t,$$

such that the cointegrating residuals v_t are fractionally integrated of order $I(d-b)$ for some $0 < b \leq d$. Obviously, for the model in equations (3.1) to (3.3), this is the case for every multiple of the vector $\left(1, -\frac{\xi_1}{\xi_2}\right)'$ and $b = \min(b_1, b_2)$.

We restrict ourselves to a bivariate set-up as is common in the literature to avoid identification problems.

According to the definition above, we thus test if among all pairs of EMU countries there exists an equilibrium relationship between the inflation rates or between IP growth rates (X_{1t} and X_{2t}), such that the persistence of deviations from the equilibrium denoted by v_t is reduced compared to that of the individual series. The degree of long memory $d-b$ in the cointegrating residual determines the persistence of deviations from the long-run equilibrium. The existence of a fractional cointegration relationship is then taken as evidence for economic integration.

In our analysis, we apply a number of semiparametric tests for the null hypothesis of no fractional cointegration. The advantage of semiparametric methods is that we do not impose any assumptions on the short-run behavior of the series, apart from mild regularity conditions. Thus, we can avoid spurious findings that might arise due to misspecification. Research on semiparametric tests for fractional cointegration has been an active field in recent years and there exist a variety of competing approaches. Whereas some approaches rely on the spectral representation of multivariate long memory processes and test whether the spectral matrix G has full rank or not, other tests are residual-based and test for the strength of integration of the cointegration residuals. To make sure that our results are robust to the way of testing, we apply tests from both strands of the literature.

Souza et al. (2018) use the fractionally differenced process $\Delta^d X_t$ and the fact that the determinant $D_{\Delta^d}(\lambda)$ of $f_{\Delta^d X}(\lambda)$ is of the form

$$D_{\Delta^d}(\lambda) \sim \tilde{G}|1 - e^{-i\lambda}|^{2b}$$

where $\tilde{G}$ is a scalar constant and $0 < \tilde{G} < \infty$. An estimate of b can therefore be obtained via a log-periodogram regression

$$logD_{\Delta^d}(\lambda) \sim \tilde{g} + 2blog|1 - e^{i\lambda}| + log\frac{\tilde{g}^*(\lambda)}{\tilde{g}}, \quad as \quad \lambda \to 0^+$$

where $lim_{\lambda \to 0^+}\ \tilde{g}^*(\lambda) = \tilde{g}^*$.

In order to make the estimation of b feasible, the empirical determinant $\widehat{D}_{\Delta^d}(\lambda)$ has to be calculated from an estimate $\widehat{f}^d_{\Delta}(\lambda)$ of the spectral density at the Fourier frequencies with $j = l, l + (2l-1), l + 2(2l-1), \ldots, m - (2l-1), m$ with $l + 1 < m < T$. The latter is obtained from the locally averaged periodogram $\widehat{f}_{\Delta^d}(\lambda_j) = \frac{1}{2l-1}\sum_{k=j-(l-1)}^{j+(l-1)} I_{\Delta^d}(\lambda_k)$, where $I_{\Delta^d}(\lambda_k$ is the periodogram of $\Delta^d X_t$. At each j the $\widehat{f}_{\Delta^d}(\lambda_j)$ estimate is thus a local average of the periodogram at frequency j and the $l-1$ frequencies to its left and right and the λ_j are spaced so that the local averages are non-overlapping.

The resulting estimator for the cointegrating strength b is given by

$$\widehat{b}_{GPH} = \left(\sum_{j=l+1}^{m} \tilde{Z}_j^{*2}\right)^{-1} \sum_{j=l+1}^{m} \tilde{Z}_j^{*} log\widehat{D}_{\Delta^d}(\lambda_j)$$

where $\tilde{Z}_j^* = Z_j^* - \bar{Z}^*$, $Z_j^* = |1 - e^{i\lambda}| = log(2 - 2cos(\lambda_j))$, and $\bar{Z}^*$ the mean of the Z_j^*. Under the null hypothesis that $b = 0$ and assuming that l and m fulfill the condition $\frac{l+1}{m} + \frac{m}{T} + \frac{1}{m} + \frac{\log m}{m} \to 0$ as $T \to \infty$, the estimate $\widehat{b}_{GPH}$ is consistent and asymptotic normal with variance $\sigma_b^2 = \frac{1}{m}(\Psi^{(1)}(2l+1) + \Psi^{(1)}(2l))$, where $\Psi^{(1)}(x) = \frac{\delta^2 log\Gamma(x)}{\delta x^2}$ is the polygamma function of order 1 and $\Gamma()$ denotes the gamma function.

The null hypothesis of no fractional cointegration can thus be tested using a simple t-test:

$$W_{SFRB} = \frac{\widehat{b}_{GPH}}{\sigma_b} \xrightarrow{d} N(0,1)$$

The method has no restrictions regarding the range of d and b and is only applicable to bivariate processes. We denote this spectral-based test in the following as SRFB18.

The test of Wang et al. (2015) is based on the sum over the fractionally differenced process $\Delta^{\hat{d}_v} X_{2t}$, which is the demeaned second component series fractionally differenced with the memory order of the potential cointegrating residual v_t. It is given by

$$W_{WWC} = T^{-1/2} \frac{\sum_{t=1}^{T} \Delta^{\hat{d}_u} X_{2t}}{\sqrt{2\pi \hat{f}_{22}(0)}}$$

where f_{22} is the spectral density of either u_{2t} or e_t in (3.1), depending on whether a triangular model or a common-components model is assumed.

Under the null hypothesis $d_v = d$ so that $\Delta^{\hat{d}_v} X_{2t}$ is $I(0)$ and the appropriately rescaled sum is asymptotically standard normal. Under the alternative $\Delta^{\hat{d}_v} X_{2t}$ is $I(b)$,so that the test statistic diverges with rate $\mathcal{O}_{\mathcal{P}}(T^b)$.

To make this test statistic feasible, the spectral density f_{22} can be estimated from the periodogram of the fractionally differenced process $\Delta^{\hat{d}_v} X_{2t}$ following the approach of Hualde (2013):

$$\widehat{f}_{22}(0) = \frac{1}{(2m+1)} \sum_{j=-m}^{m} I_{\Delta^{\hat{d}} X_2}(\lambda_j)$$

where $I_{\Delta^{\hat{d}} X_2}(\lambda_j)$ is the periodogram of $\Delta^{\hat{d}} X_2(\lambda_j)$.

While Wang et al. (2015) are agnostic about the method that is used for the estimation of the memory parameters d and d_v, they assume that $d > 1/2$ so that the cointegrating vector can be estimated using ordinary least squares. The memory orders can be estimated from $\widehat{v}_t^{OLS}$ and X_{2t} using any of the common semiparametric estimates such as ELW with bandwidth m as in $\widehat{f_{22}}$ that fulfills the usual bandwidth conditions.

The method does not impose any restrictions on the fractional cointegrating strength b. According to the Monte Carlo simulations of Leschinski et al. (2020), the non-stationarity requirement ($d > 1/2$) can be circumvented if the cointegrating residual v_t is based on the NBLS estimate of the cointegrating vector instead of the OLS estimate.

In contrast, the test of Chen and Hurvich (2006) (denoted by CH06) is directly based on $\hat{d}_v$, but the cointegrating space is estimated by the eigenvectors of the averaged and tapered periodogram matrix local to the origin. The process X_t is assumed to be stationary after taking $(q-1)$ integer differences which allows $d \in (q-1.5, q-0.5)$. In order to account for possible over-differentiation the complex-valued taper $h_t = 0.5(1-e^{i2\pi t/T})$ of Hurvich and Chen(2000) is applied to the data. The tapered discrete Fourier transform (DFT) and periodogram of X_t are defined by

$$
\begin{aligned}
w_X^{tap}(\lambda_j) &= \frac{1}{\sqrt{2\pi \sum_t |h_t^{(q-1)}|^2}} \sum_{t=1}^{T} h_t^{(q-1)} X_t e^{i\lambda_j t} \\
I_X^{tap}(\lambda_j) &= w_X(\lambda_j)\overline{w_X(\lambda_j)}
\end{aligned}
$$

Based on the tapered periodogram, define the averaged periodogram matrix of X_t by $I_X^{av}(\lambda_j) = \sum_{j=1}^{m_3} Re(I_X^{tap}(\lambda_j))$, where m_3 is a fixed positive integer fulfilling $m_3 > p+3$. The eigenvalues of $I_X^{av}(\lambda_j)$ sorted in descending order are denoted by $\hat{\delta}_{\alpha,I_X^{av}}$ and the corresponding eigenvectors are given by $\hat{\chi}_{\alpha,I_X^{av}}$ for $\alpha = 1, \ldots, p$. Under the alternative hypothesis, if there are $r > 0$ cointegrating relationships, the matrix consisting of the first r eigenvectors provides a consistent estimate of the cointegrating subspace. To construct a test for the null hypothesis of no fractional cointegration, the potential cointegrating residuals v_t are estimated by multiplying X_t with the eigenvectors $\hat{\chi}_{\alpha,I_X^{av}}$ so that $\hat{v}_{at}^{av} = \hat{\chi}'_{\alpha,I_X^{av}} X_t$, for $\alpha = 1, \ldots, p$. The memory of the p residual processes is estimated with the local Whittle estimator using bandwidth m, but calculated using shifted Fourier frequencies $\lambda_{\tilde{j}}$ with $\tilde{j} = j + (q-1)/2$ to account for the tapering of order q. These estimates are denoted by $\hat{d}_{v_\alpha,\widetilde{LW}}$ and they remain consistent and asymptotic normal. Since there can be at most $p-1$ cointegrating relationships in a $p-$dimensional time series, the first residual corresponding to the largest eigenvalue cannot be a cointegrating residual. Its memory must therefore equal the common memory d of X_t . In contrast, the last residual $\widehat{v}_{pt}^{av}$ corresponding to the smallest eigenvalue is most likely to be a cointegrating residual if there is cointegration so that its memory is reduced by b under cointegration.

The test idea of Chen et al. (2006) is therefore to compare the estimated memory orders from the residual series $\widehat{v}_{1t}^{av}$ and $\widehat{v}_{pt}^{av}$ that correspond to $\widehat{d}$ (first residual) and $\widehat{d}_u$(last residual). Chen et al. (2006) show that

$$\sqrt{m}(\widehat{d}_{v_\alpha,\widetilde{LW}} - \widehat{d}_{v_b,\widetilde{LW}}) \xrightarrow{d} N\left(0, V_{CH,q}\left(1 - \frac{G_{ab}^2}{G_{aa}G_{bb}}\right)\right)$$

with $V_{CH,q} = \frac{1}{2}\frac{\Gamma(4q-3)\Gamma^4(q)}{\Gamma^4(2q-1)}$. A conservative test statistic is therefore given by

$$W_{CH} = \sqrt{m}\frac{\left(\widehat{d}_{v_1,\widetilde{LW}} - \widehat{d}_{v_p,\widetilde{LW}}\right)}{\sqrt{V_{CH,q}}}$$

The test rejects if W_{CH} is larger than the standard normal quantile $z_1 - \alpha/2$. It is very versatile, since it does not impose restrictions on the cointegration strengh b and can be applied to stationary as well as non-stationary long-memory processes, but it requires a priori knowledge about the location of d in the parameter space to determine the order of differencing. Obviously, these two tests are residual-based.

Leschinski et al. (2020) suggest that the three testing procedures have better performance among a group of eight semiparametric tests, particularly when testing for fractional cointegration. In bivariate cases, SRFB18 presents the best performance.

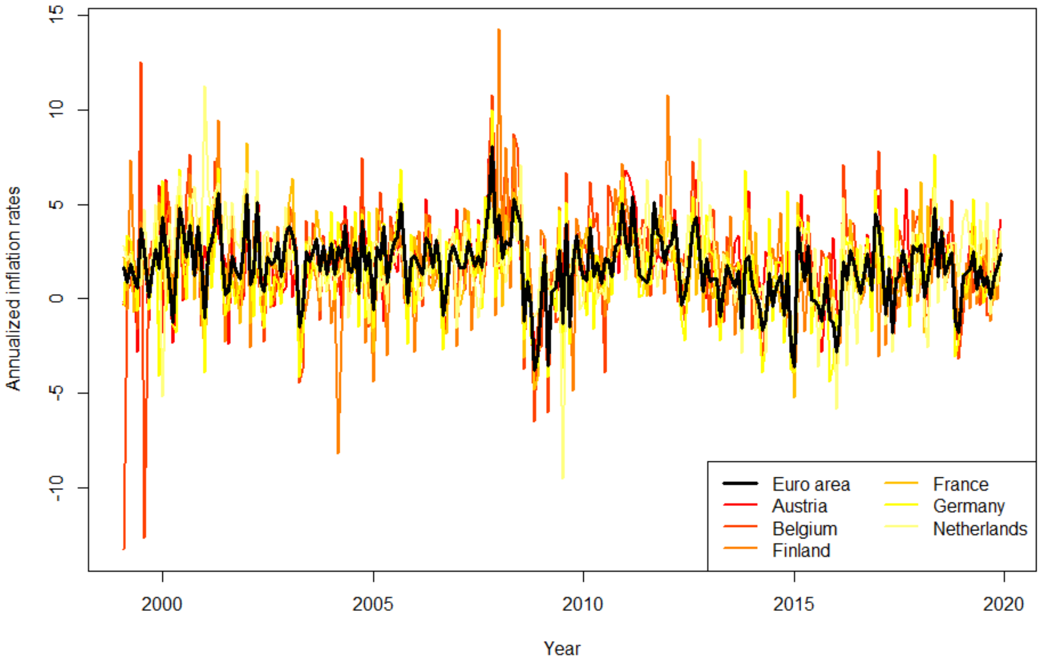
3.3 Data

We conduct the analysis of macroeconomic convergence for 11 EMU countries: Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Netherlands, Portugal and Spain. As is common in the literature, we term the group of Austria, Belgium, Finland, France, Germany and the Netherlands the “core countries” of EMU, while we call the group of Greece, Italy, Ireland, Spain and Portugal the “periphery countries”. The data sample ranges from 1999:1 to 2019:12. Our sample thus starts with the official start of the EMU with the ECB acting as single central bank for the monetary union.

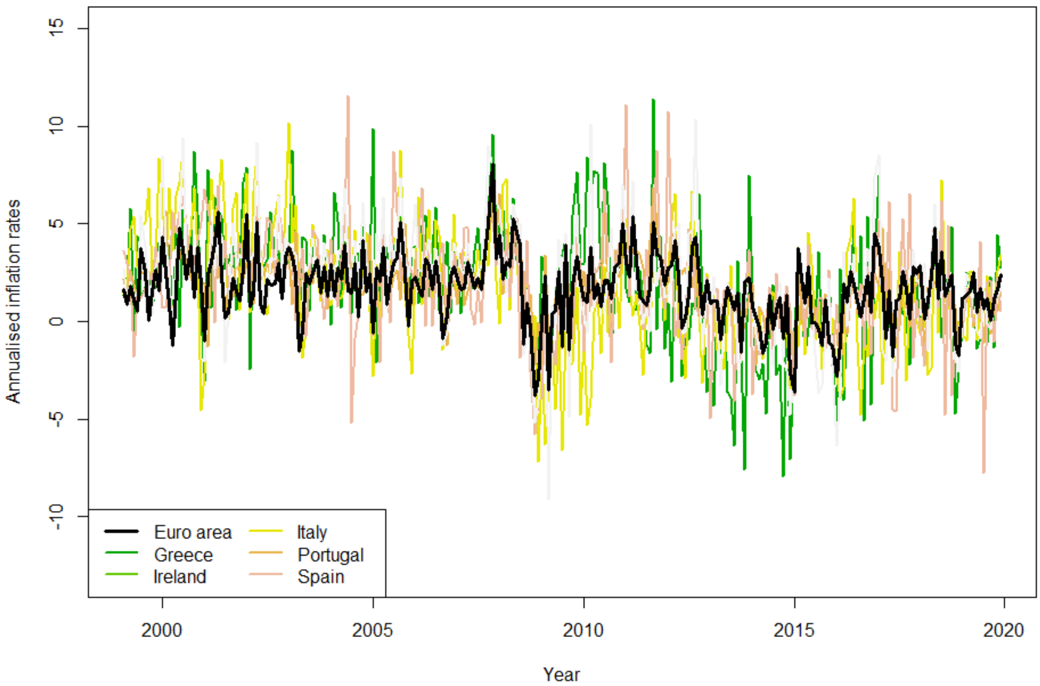
Data for inflation is obtained from Eurostat and measured with the monthly seasonally unadjusted Harmonized Indices of Consumer Prices for all items, with 2015 as base year (2015 = 100). We seasonally adjust each series by using the X-13 R package, developed by the United States Census Bureau. Then, we define annualized inflation rates for each country i as

$$\pi_{it} = 1200(log(HCPI_{it}) - log(HCPI_{it-1}))$$

The series are shown for all countries in our sample in Figure 3.1. Visual inspection suggests that all EMU countries in our sample experienced a drop in inflation at the start of the financial crisis in 2008, with somewhat more volatile inflation rates after 2008. Overall, inflation rates in the periphery countries seem more volatile than those in the core countries.

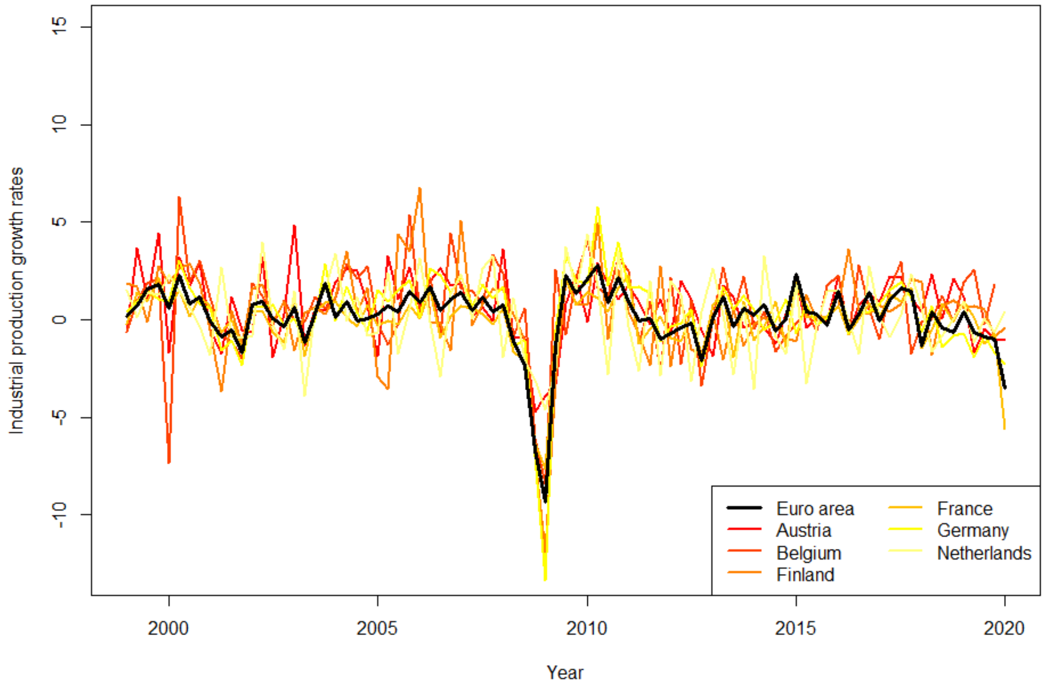


(a) Annualized inflation rates in core countries

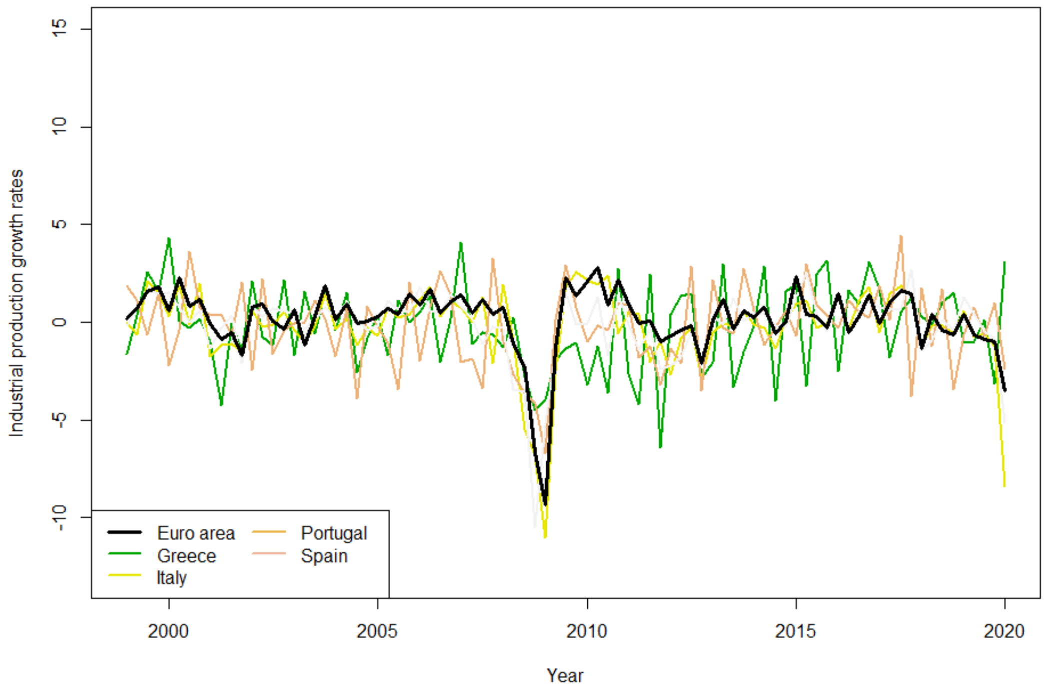


(b) Annualized inflation rates in periphery countries

Figure 3.1: Annualized inflation rates of EMU countries.



(a) Industrial production growth rates in core countries



(b) Industrial production growth rates in periphery countries

Figure 3.2: Industrial production growth rates of EMU countries.

Data for quarterly seasonally adjusted industrial production growth rates are obtained from the Federal Reserve Economic Database (FRED) by the St. Louis Fed. The data is shown in Figure 3.2. As in the case of inflation rates, the pronounced drop in industrial production in 2008 is clearly visible in all countries of our sample. Before the crisis, countries like France, Germany and Spain experienced very stable IP growth, while countries like Finland, Greece and Portugal exhibit higher IP volatility. After the financial crisis, the negative effects of the European sovereign debt crisis on industrial production growth rates are clearly observable in Greece and Portugal. Here, we exclude Ireland from our sample because of data irregularities.

3.4 Analysis of Convergence and Divergence of EMU Inflation Rates

3.4.1 Measuring the degree of long memory in EMU inflation rates

In this section, following Kumar and Okimoto (2007), we model inflation rates as a fractionally integrated process and estimate the memory parameters as a measure of persistence for the largest economies in the EMU. Memory parameters for each country are shown in Table 3.1, where we differentiate between the so-called "core" economies of Austria, Belgium, Finland, France, Germany and the Netherlands, and the so-called "periphery" countries Greece, Ireland, Italy, Portugal and Spain. For our estimation, we use the exact local Whittle estimator of Shimotsu (2010) with a 0.75 bandwidth d-value ($m = T^{0.75}$). As a direct extension of Shimotsu et al. (2005), this estimator has the advantage of allowing for non-zero means, while the properties of consistency and asymptotically normal distribution for all values of d continue to hold. The estimator is given by:

$$\hat{d}_{ELW} = argmin_{-1<d<3.5} \left\{ log\hat{G}_m(d) - d(\frac{2}{m}\sum_{j=1}^{m} log\lambda_j) \right\}$$

where $\lambda_j = 2\pi j/T$, $\hat{G}_m(d) = m^{-1}\sum_{j=1}^{m} I_{\Delta^d x}(\lambda_j)$, and $I_{\Delta^d x}(\lambda)$ denotes the periodogram of the fractionally differenced process $(1-L)^d(X_t)$. Similar to other findings in the literature Hualde and Iacone (2017a); Meller and Nautz (2012), inflation persistence is higher for the periphery countries. Overall, the memory parameters are ranging from 0.26 to 0.44 in periphery countries and from 0.12 to 0.33 in core countries. In fact, the mean values of the order of fractional integration are 0.24 and 0.34 for core and periphery countries, respectively.

Table 3.1: Memory estimates for inflation rates

Austria	Belgium	Finland	France	Germany	Netherlands	Mean
0.23	0.24	0.3	0.21	0.12	0.33	0.24
Greece	Ireland	Italy	Portugal	Spain		Mean
0.35	0.36	0.44	0.28	0.26		0.34

Note: Exact local Whittle estimates of d with bandwidth $m = T^{0.75}$.

3.4.2 Testing for fractional cointegration in EMU inflation rates

Following Leschinski et al. (2020) findings, we perform all tests at the 5% significance level. The bandwidth is selected as $m = T^{0.75}$ for all three testing procedures.[1] The trimming parameter r is set to 3 for SRFB18 and the integer for averaging the periodogram is 25 for CH06. As fractional cointegration needs as a core assumption that the order of integration is equal between the series, we use pairwise tests as suggested by Robinson and Yajima (2002) to test for the equality of the memory parameters. The results shown in Table 3.2 suggest that the hypothesis of a common memory parameter is rejected for the case of Germany with Greece, Ireland and Italy, as well as for the pair of France-Italy at the 5% significance level. Therefore, these pairs are not included in our analysis. Next, we test for pairwise cointegrating relationships under the null hypothesis of no fractional cointegration.

An example of our analysis is in Table 3.9 in the appendix A where we report the test

[1] We also use two different bandwidths, $m = T^{0.65}$ and $m = T^{0.7}$. Our main findings are similar for both cases.

Table 3.2: Pair-wise testing for common memory integration in inflation rates

EMU countries	Belgium	Finland	France	Germany	Netherlands	Greece	Ireland	Italy	Portugal	Spain
Austria	0.11	0.65	0.08	0.99	0.93	1.10	1.21	1.89	0.55	0.48
Belgium		0.59	0.2	1.04	0.8	1.06	1.16	1.77	0.43	0.36
Finland			0.77	1.58	0.28	0.43	0.60	1.22	0.1	0.25
France				0.93	0.97	1.26	1.38	2.15**	0.63	0.58
Germany					1.78	1.98**	2.22**	2.77**	1.39	1.42
Netherlands						0.15	0.26	0.81	0.4	0.53
Greece							0.11	0.75	0.56	0.76
Ireland								0.61	0.68	0.86
Italy									1.35	1.61
Portugal										0.14

Note: Results of Robinson and Yajima (2002) test. ** denotes significance of the test statistic at the 5% level where the critical value is 1.9599.

results for Germany as the reference country, since it is the largest economy within the EMU. There is evidence for fractional cointegration of inflation rates with the German counterpart in all seven countries we were allowed to test. For the case of Austria, the null hypothesis is rejected for all three testing procedures. France, Netherlands and Spain suggest inflation convergence only for the WWC15 test. We repeat the testing procedures for all possible pairs.

As a second step, we estimate the cointegration vector β as well as the memory parameters d_v by using Narrow Band Least Squares estimation. Robinson (1994) shows that NBLS estimation is consistent under stationary cointegration whereas an OLS approach might not retain consistency.

Table 3.3: Pair-wise memory estimates on the cointegrating residuals of inflation rates

EMU countries	Austria	Belgium	Finland	France	Germany	Netherlands	Greece	Italy	Portugal	Spain
Austria		**0.04**		**0.12**	**0.03**					
Belgium	**0.18**			**0.18**	**0.19**			**0.16**	**0.1**	**0.15**
Finland						**0.24**		**0.22**		**0.23**
France	**0.14**	**0.1**			**0.1**					
Germany	**-0.06**	**0.03**		**0.01**						
Netherlands			**0.27**					**0.28**	**0.19**	**0.24**
Greece									**0.26**	**0.24**
Italy		**0.13**	**0.23**			**0.29**			**0.09**	**0.12**
Portugal		**0.2**				**0.23**	**0.12**	**0.21**		**0.05**
Spain		**0.21**	**0.24**			**0.24**	**0.11**	**0.2**	**0.11**	

Note: Exact Local Whittle estimates, d_v, with a $m = T^{0.75}$ bandwidth based on the Narrow Band Least Squares estimation of the cointegration vector β.

We thus repeat the NBLS estimation between all EMU country pairs and we present the estimated values of the memory of the residuals in Table 3.3. Only the cases where the hypothesis of a common memory parameter is not rejected, at least one of the testing procedures are statistically significant and the memory of d_v is less than the memory of both of the individual series are presented (e.g. Ireland is not cointegrated with any of the countries, therefore is not presented in the following table). As can be seen in Table 3.3, results are not symmetric as it is expected for cointegration relations. This is due to the irregular behavior of the inflation series during the crisis period. It is well known that such data irregularities can lead to asymmetric cointegration (see for example Enders and Siklos (2001)).

For the core countries, inflation rates between Germany, Austria and the rest of the countries are estimated to have the strongest cointegration relation, as the adjustment to equilibrium is achieved faster after potential shocks (d_v ranges between -0.06 and 0.04). Belgium, France, Finland and the Netherlands are all estimated to have fractional cointegration relationships, but with slower adjustment to equilibrium overall.

For the periphery countries (see the lower part of Table 3.3), the memory parameters that belong to the pairs of Italy, Portugal and Spain are lower than the parameters among the core-periphery or periphery-core pairs. On average, we observe that cointegration pairs cluster among core and periphery countries, with fewer mixed core-periphery cointegration pairs. This suggests that over the whole estimation period, separate 'convergence clubs' for EMU inflation rates may be observed. This could be worrisome as it might be indicative of macroeconomic divergence processes, which could lead to current account disequilibria.

3.4.3 Testing for a break in fractional integration of EMU inflation rates

In addition to our analysis of fractional cointegration, we test for a break in the order of fractional integration in inflation rates. This allows us to test whether persistence and thereby potential convergence or divergence processes in inflation changed over time. This could be due, for instance, to the disruptions of the global financial crisis and the European sovereign debt crisis. Hence, we perform a regression-based Lagrange Multiplier test introduced by Martins and Rodrigues (2014) that generalizes the conventional integration approaches to the fractional integrated process context. Their method is based on the recursive forward and reverse estimation of the Hassler and Breitung (2002) and Robinson (1994) test statistics. The X_t series is first filtered as in (3.1-3.3) and persistence changes occur in the interval $[\tau T]$, with $\tau \in [\Lambda_l, \Lambda_u]$ and $0 < \Lambda_l < \Lambda_u < 1$ as is usually assumed in the structural breaks literature. Then, the auxiliary regression is given by:

$$x_t = \phi(\tau) x^*_{t-1} + e_t, t = 2, \ldots, [\tau T] \tag{3.5}$$

where $x_t = (1-L)^{d_0} y_t$ and $x^*_{t-1} = \sum_{j=1}^{t-1} \frac{x_{t-j}}{j}$. The test statistic is formed by recursively estimating (3.5), the least squares t-statistic for $\hat{\phi}(\tau) = 0$ and the auxiliary regression in the time-reversed series for the remaining $(1-\tau)T$ observations. This method is able to work with unknown date of the shifts, trends and serial correlations.

In what follows we will denote by $\hat{s}$ the estimated point of the persistence shift and by d_1 the order of integration before the shift and by d_2 the order of integration after the break. Table 3.4 suggests evidence of a structural break in inflation persistence for all EMU countries. Specifically, Austria, Belgium, Finland, France, Greece, Ireland and Italy's estimated breakpoints ($\hat{s}$) occur from October, 2006 to October, 2008. This corresponds to the time period covering the beginning of the subprime crisis, i.e. a heightened level of financial stress, to the beginning of the global financial crisis following the default of Lehman Brothers on September 15, 2008. Comparing d_1 and d_2, the results suggest a shift from $d \leq 0$ to $d > 0$ at the time of $\hat{s}$, that is an increase in inflation persistence after the break. The alternative hypothesis of decreasing memory holds for a smaller number of countries. For Finland, Greece, Ireland, Italy, Portugal and Netherlands we find (additional) memory shifts between April, 2014 and October, 2016, with inflation persistence returning to its pre-crisis lower levels.

In summary, we find that inflation persistence since the beginning of EMU is stationary mean-reverting. A test in persistence change is showing evidence of breaks in the beginning of the financial crisis where we observe memory parameters around zero in the pre-crisis period for the majority of the sample, and substantially higher values

Table 3.4: Persistence change test in inflation rates

EMU country	Test-statistic	Date	d_1	d_2
Austria	-3.84***	10/2006	-0.25	0.27
Belgium	-2.68***	09/2007	-0.01	0.35
Finland	-1.83**	10/2007	0.02	0.34
	-2.67*	02/2016	0.28	-0.18
France	-1.84*	07/ 2007	-0.21	0.39
Germany	-2.92***	06/2005	-0.09	0.17
Netherlands	-1.9317*	12/2014	0.35	0.10
Greece	-2.96***	05/2007	-0.19	0.37
	-3.87***	10/2014	0.36	-0.1
Ireland	-1.98**	10/2008	0.13	0.32
	-2.60**	10/2016	0.40	-0.15
Italy	-4.79***	08/2007	-0.02	0.45
	-2.28*	12/2014	0.46	0.14
Portugal	-3.82***	04/2014	0.32	-0.15
Spain	-1.75*	06/2005	-0.06	0.33

Note: Results of Martins and Rodrigues (2014) test where *, **, *** indicate levels of significance at 10%, 5%, 1%, respectively. "d_1" and "d_2" refer to the memory parameters of the respective subperiod of the estimated breakpoint.

around 0.4 in the post-crisis period. Moreover, the inflation persistence during the crisis period increased in particular in the periphery countries, implying potentials for inflation divergence. Finally, while we find evidence of fractional cointegration in inflation among the full sample, the core countries tend to be fractionally cointegrated with a faster speed of adjustment, see also Hualde and Iacone (2017b) and Karanasos et al. (2016).

3.5 Analysis of Convergence and Divergence of EMU Industrial Production

3.5.1 Measuring the degree of long memory of EMU industrial production

Following the analysis of convergence in EMU inflation rates in the previous section, we next investigate the convergence of industrial production rates between the same countries, except Ireland. First, as in subsection 3.4.1 we estimate the memory parameter of each country. Results are presented in Table 3.5.

Table 3.5: Memory estimates for industrial production growth rates

Austria	Belgium	Finland	France	Germany	Netherlands	Mean
0.31	0.08	0.27	0.39	0.47	-0.12	0.23
Greece	Italy	Portugal	Spain			Mean
0.11	0.68	0.22	0.56			0.39

Note: Exact local Whittle estimates of d with bandwidth $m = T^{0.75}$.

We observe that there are more cases of memory estimates with $d > 0.5$ in the periphery countries (Italy and Spain), while in the core countries almost all memory estimates are stationary mean-reverting with $-0.5 < d < 0.5$. In line with our previous results for inflation, we find that long memory parameters of industrial production are on average higher in the periphery countries of EMU.

3.5.2 Testing for fractional cointegration of EMU industrial production

Next, we test for fractional cointegration in EMU industrial production among all possible country pairs. In Table 3.6, the hypothesis of memory equality is rejected for almost all cases in Italy and the Netherlands. Therefore, they are excluded from our analysis, as before. Similarly to the previous section, we present in Table 3.10 in the appendix A the test results for cointegration between industrial production in Germany and the remaining EMU countries.

Table 3.7 reports the NBLS estimates of the memory parameters, d_v, of the IP fractional cointegration relation in all country-pairs. Here, asymmetric cointegration for industrial production country-pairs is even more pronounced than in the case of inflation. This is again due to the drop of industrial production during the financial crisis. We also exclude the Netherlands since it is not fractionally cointegrated with any of the remaining countries. Comparing our results for fractional integration of EMU inflation and IP, inflation convergence was suggested mostly between core and between periph-

Table 3.6: Pair-wise testing for common memory integration in industrial production growth rates

EMU countries	Belgium	Finland	France	Germany	Netherlands	Greece	Italy	Portugal	Spain
Austria	1.208	0.041	0.389	0.742	1.965**	0.475	1.972**	0.510	1.356
Belgium		1.231	1.562	1.9496	0.9133	0.583	2.978**	0.577	2.334
Finland			0.456	0.784	2.094**	0.444	2.033**	0.453	1.380
France				0.415	2.420**	0.834	2.058**	0.929	1.224
Germany					2.827**	1.091	1.652	1.259	0.798
Netherlands						1.422	3.729**	1.348	3.089**
Greece							2.285**	0.032	1.761
Italy								2.649**	0.776
Portugal									2.008**

Note: Results of Robinson and Yajima (2002) test. ** denotes significance of the test statistic at the 5% level where the critical value is 1.9599.

ery countries, whereas convergence in IP growth rates appears stronger between mixed groups of core/periphery countries. For instance, German IP is fractionally cointegrated with Austria and Finland, but also Italy and Spain. This result could mirror the strong trade links between these countries. Overall, it appears that nominal convergence does not necessarily imply real convergence and *vice versa*. The fact that we observe more bilateral cointegration relationships among core or between core and periphery countries, but few among periphery countries, suggests that convergence in industrial production might be driven by the core countries.

Table 3.7: Pair-wise memory estimates on the cointegrating residuals of industrial production growth rates

EMU countries	Austria	Belgium	Finland	France	Germany	Greece	Italy	Portugal	Spain
Austria		**-0.07**	**-0.26**	**0.04**	**0.06**	**0.06**			
Belgium	**-0.12**		**-0.19**	**-0.03**					
Finland	**0.09**	**0.06**		**0.18**	**0.09**				**0.08**
France	**-0.02**	**0.00**	**0.09**			**0.09**		**0.02**	
Germany	**0.29**		**0.15**				**0.21**		**0.32**
Greece	**0.05**			**0.07**				**0.07**	
Italy					**0.2**				**0.12**
Portugal				**0.07**		**0.09**			
Spain			**0.04**		**0.28**		**0.07**		

Note: Exact Local Whittle estimates, d_v, with a $m = T^{0.75}$ bandwidth based on the Narrow Band Least Squares estimation of the cointegration vector β.

Table 3.8: Persistence change test in industrial production growth rates

EMU country	Test-statistic	Date	d_1	d_2
Austria	-2.57***	2005.03	-0.18	0.37
Belgium	-2.13**	2004.04	-0.5	0.16
France	-2.06**	2008.03	0.19	0.31
	-2.56***	2011.01	0.59	-0.4
Italy	-2.53***	2008.03	0.31	0.82
Portugal	-2.05**	2008.02	-0.32	0.43
Spain	-3.69***	2007.01	0.38	0.58
	-4.98***	2009.03	0.7	0.37

Note: Results of Martins and Rodrigues (2014) test where *, **, *** indicate levels of significance at 10%, 5%, 1%, respectively. "d_1" and "d_2" refer to the memory parameters of the respective subperiod of the estimated breakpoint.

3.5.3 Testing for a break in fractional integration of EMU industrial production

Finally, we apply the test by Martins and Rodrigues (2014) to test for a break in the order of fractional integration in EMU industrial production. The results are shown in Table 3.8 . As can be seen, we find a significant break in IP persistence in most, but not all countries. In France, Italy, Portugal and Spain the test suggests a significant breakpoint in presistence of industrial production around the beginning of the global financial crisis, with higher persistence in the post-crisis period. This suggests that the large drop in IP observed in all countries of our sample led to changes in the long memory of the series in these countries. However, for France and Spain we find a second breakpoint in 2011 and 2009, respectively. In these two countries, after the second break estimated persistence is at pre-crisis levels or even lower, suggesting that the change in persistence during the crisis was only temporary in these cases.

3.6 Conclusions

In this paper, we apply methods of fractional cointegration to investigate the degree of real and nominal convergence between EMU countries. Specifically, we model both inflation rates and industrial production growth rates as fractionally integrated processes and estimate the memory parameters as a measure of persistence. The analysis covers the full EMU period from January, 1999, to December, 2019 for 11 EMU countries consisting of both "core" and "periphery" countries. Moreover, we test for breaks in persistence with the test by Martins and Rodrigues (2014).

To test for the existence of fractional cointegration relationships for inflation and IP among all country pairs of our sample, we estimate the memory parameters of the residuals d_v of the fractional cointegration relation using the NBLS estimation on the cointegrating vector β. We find evidence of fractional cointegration among EMU countries for both inflation and real IP growth. In the case of inflation, the results suggest more cointegration relations among core or among periphery countries. This implies the existence of separate "convergence clubs" within EMU inflation rates over the whole estimation period. This could be an indicator of macroeconomic divergence processes. In the case of EMU industrial production growth rates, we find evidence of fractional cointegration either among core or among mixed core-periphery country pairs. This could be indicative of real convergence processes being driven by trade links among the EMU countries. Overall, we find that nominal convergence does not necessarily imply real convergence and *vice versa*.

In addition, our results suggest breaks in the persistence of both inflation and industrial production growth rates around the beginning of the financial crisis in 2007-08 for a majority of countries in our sample. For some countries, we find a second break in 2009-11 (IP) or 2014-16 (inflation) marking the end of the crisis period. Generally, we find lower persistence of inflation and IP growth rates before the crisis and substantially higher persistence in the crisis period. In case of a second break, persistence reverts back to pre-crisis levels. The higher persistence of the series during the crisis period carries the danger of diverging processes in case of asymmetric shocks.

To sum up, our analysis gives a detailed picture of time-variation in real and nominal convergence processes since the start of the EMU. While the financial crisis and the European sovereign debt crisis was a period of potential divergence in both inflation and industrial production, there is nevertheless some evidence that a) the crisis period ended in some cases before the end of our sample and b) we still find evidence of fractional cointegration in many country-pairs for the full sample period. Still, in light of the current crisis due to the COVID-19 pandemic, our results tell a cautious tale about the potential vulnerability of macroeconomic convergence within EMU during economic crisis periods.

The results suggest several implications for macroeconomic policies within the EMU: First, we do not find evidence of stable long-run relationships of inflation and industrial production in all country-pairs. This implies that macroeconomic convergence within EMU is far from complete. The fact that convergence in inflation does not necessarily imply convergence in real IP (and *vice versa*) suggests that macroeconomic convergence is affected by complex interrelations between trade effects, domestic shocks and economic structures and macroeconomic policy. Hence, a high degree of cooperation in economic policy will be needed in order to push real macroeconomic convergence further. Second, there is evidence for separate "convergence clubs" within EMU regarding inflation both before and after the financial crisis of 2008. This implies a continuing threat of real appreciation or depreciation among countries in the currency union and potentially inefficiency of monetary policy if the inflation rates deviate in levels. Thus, again a need for further country-specific policy measures to foster convergence also in terms of inflation is highlighted by our results. Third, while the financial and economic crisis period shows higher persistence in real and nominal processes, such that negative deviations may hold for longer periods, our results suggest that in some countries the increase in persistence was reversed after some time. Hence, while there is still a lot of scope for further economic integration within EMU, the good news is that at least in some cases the additional adverse effects of the crisis period seem to abate after some years.

Appendix A

The results of the three semiparametric tests we used in our analysis of pair-wise fractional cointegration, with Germany as a reference country, are presented in the next tables. The tests by Souza et al. (2018),Wang et al. (2015),Chen et al. (2006) are denoted as SRFB18, WWC15 and CH06 respectively. Here, the null hypothesis is no fractional cointegration.

Table 3.9: Pair-wise testing for stationary fractional cointegration in inflation rates

Germany/	SRFB18	WWC15	CH06
Austria	2.2855**	23.355**	2.4783**
Belgium	1.032	7.368**	1.7854**
Finland	1.3497	9.0378**	1.3954**
France	1.5783	10.9783**	1.0047
Netherlands	0.056	9.7593**	1.0851
Portugal	0.9944	10.1987**	1.9513**
Spain	-0.6128	7.8308**	1.3307

Note: Critical values at $\alpha = 5\%$ are 1.960 for both SRFB18 and WWC15, as well as 1.386 for CH06. ** denotes significance of the test statistic at the 5% level. Here, H_0: no fractional cointegration.

Table 3.10: Pair-wise testing for stationary fractional cointegration in industrial production growth rates

Germany/	SRFB18	WWC15	CH06
Austria	5.382**	3.377**	2.253**
Belgium	2.004**	1.37	2.205**
Finland	3.760**	0.505	1.310
France	4.448**	0.001	1.264
Greece	2.301**	0.348	1.091
Italy	5.884**	0.42	1.744**
Portugal	2.654**	0.13	0.932
Spain	3.398**	0.191	1.007

Note: Critical values at $\alpha = 5\%$ are 1.960 for both SRFB18 and WWC15, as well as 1.386 for CH06. ** indicates significance at the 5% level. Here, H_0: no fractional cointegration.

Chapter 4

The spillover effect of ECB's policy rate on Euro Area economies: A TVP-FAVAR approach

4.1 Introduction

Over the last two decades the Economic and Monetary Union (EMU) of the European Union is operating with a common currency and a single monetary policy under the authority of the European Central Bank (ECB). ECB's primary objective is to ensure price stability along with balanced economic growth for the country members. However, during and after the financial crisis followed by the European sovereign debt crisis, nominal interest rates declined globally. Specifically, ECB lowered its policy rate initially from 4.25 to 1.00, then raising to 1.5 and finally falling to zero lower bound (ZLB), right after the European debt crisis (Hartmann and Smets (2018), Laine (2020)).

As a significant number of European countries have faced debt crises during the last years and have employed several fiscal measures in order to shrink their deficits immediately. Examples include Italy, Greece, Spain, Portugal, etc. In this research, we aim to investigate the impact of a common monetary policy shock on the Euro Area(EA) country members. Given their diversity in economic and financial structure, different results are anticipated both in terms of size and timing. There is a large and growing literature on the effects of monetary policy shocks on the countries of the EA (Barigozzi et al. (2014), Blaes (2009), Boivin et al. (2008), Galariotis et al. (2018), Peersman (2004), Potjagailo (2017)) focusing on the asymmetric responses both on aggregate and country levels.

Vector autoregressions (VAR) are typically used in the examination of unanticipated policy shocks developed by Sims (1980). Nevertheless, in our analysis we are taking advantage of recent developments in time series analysis, employing time-series econometric techniques that have been developed during the last 15 years and have not been used extensively in the empirical literature. More precisely, we aim in using models that combine the standard VAR analysis with factor models. These models are referred to as factor-augmented vector autoregressive models (FAVAR) in the relevant literature and have been developed by Bernanke et al. (2005). The dynamic factor models are able to summarize the information from a large number of macroeconomic and financial time series into a small number of unobserved factors, which capture the common variation or co-movements of the initial large number of available economic series. These factors will increase the amount of information in the VAR and the innovations will span the space of the structural disturbances. In other words, FAVARs are expected to mitigate the omitted variable bias that is often inherent in standard small-dimensional 2 VAR models.

Barigozzi et al. (2014), Boivin et al. (2008) and Laine (2020) are using the FAVAR methodology to investigate the effects of monetary policy across the Euro area countries. Barigozzi et al. (2014) dataset includes 10 countries between 1983q1 and 2007q4 for 237 series and their results show asymmetries in the transmission between Northern and Southern countries for prices and unemployment whereas no difference appears in the case of output. These results are explained by the lack in both of price flexibility and market competition. Similarly, Boivin et al. (2008) within the same time period (1980-2007) are employing 245 quarterly series for the six largest EA economies by dividing in pre- and post- EA periods. Their results provide useful insights of the transmission mechanism as they show a greater homogeneity across countries, and an overall reduction in the effects of monetary shocks during the post-EA period. Laine (2020) are examining the ECB's conventional monetary policy on the real economy. The dataset includes 90 series from January 1999 to July 2017. The results suggest that the transmission to real economy has weakened after the financial crisis of 2008, as well as the importance of including time-varying parameters in FAVAR models as the responses of economic variables vary over time.

We contribute to the existing literature by using a FAVAR model with time varying coefficients and stochastic volatility. In that way, we focus on a large number of variables and investigate simultaneously the effects in size and over time. Also, we study the effects not only on a aggregate level where the effects can cancel out over time but also on each EA country member. Thus, our study extends the previous studies as we allow the FAVAR model parameters to change over time and by employing a large dataset that spans from January, 2000 until December, 2019, we take into account the recent period of economic turmoil following the financial crisis in 2008. We strengthen the evidence of asymmetric transmission for the cases of inflation, unemployment rates and long-term government bonds rates between core and periphery countries. We further compare our results with a TVP-VAR model and we show that TVP-FAVAR can capture better monetary policy's effectiveness. We find that even though the lower interest rates in the post crisis period, industrial production, prices and long-term government bonds presented more persistent reactions to a potential monetary shock.

The rest of the paper is organized as follows. In Section 2 the TVP-FAVAR methodology is presented. Sections 3 describes the data selection for the analysis. In Section 4 the empirical results are presented, while Section 5 concludes.

4.2 Methodology

4.2.1 The TVP-FAVAR model

In general, a TVP-FAVAR can initially be described as a factor model for multivariate time series with drifting coefficients and stochastic volatility. We therefore consider the factor equation as in Molteni and Pappa (2017)

$$X_t = \Lambda^f F_t + \Lambda^r R_t + u_t \tag{4.1}$$

where X_t is a $N \times 1$ vector of macroeconomic and financial time series and F_t a $K \times 1$ unobserved factor that represents the potential forces on economy incorporated in X_t. R_t is a $M \times 1$ observed factor that represents the policy instrument ($N >> M$). Λ^f and Λ^r are the matrices of factor loadings of dimensions $(N \times K)$ and $(N \times M)$, respectively. Both factor loading cases are associating F_t and R_t with X_t. Last, u_t is a $N \times 1$ vector of error terms, $u_t \sim N(0, I_N)$. Error terms u_t and both factors (F_t, R_t) are assumed to be uncorrelated at all leads and lags, as well as mutually uncorrelated.

$$E(u_{i,t}F_t) = E(u_{i,t}R_t) = E(u_{i,t}u_{j,s}) = 0$$

for all $i, j = 1, \dots, n \wedge t, s = 1, \dots, t$ and $i \neq j \wedge t \neq s$.

In order to assess the monetary policy actions, the TVP-FAVAR now takes the following form:

$$y_t = b_{1,t}y_{t-1} + \cdots + b_{p,t}y_{t-p} + \upsilon_t \tag{4.2}$$

where $y_t = [F_t, R_t]$ is a $((q = m + k) \times 1)$ vector of both observed and unobserved factors, $B_{i,t}$, $i = 1, \dots, p$ are the $q \times q$ matrices of time varying coefficients and $\upsilon_t \sim N(0, \Omega_t)$.

We define the decomposition of the Ω_t covariance matrix as in Primiceri (2005)

$$\Omega_t = A_t^{-1}\Sigma_t\Sigma_t^{'}(A_t^{-1})' \tag{4.3}$$

where $\Sigma_t = diag(\sigma_{1,t}, \ldots, \sigma_{q,t})$ and A_t is the lower triangular matrix

$$A_t = \begin{bmatrix} 1 & 0 & \ldots & 0 \\ \alpha_{21,t} & 1 & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ \alpha_{q1,t} & \ldots & \alpha_{qq-1,t} & 1 \end{bmatrix}$$

The time varying parameter are modelled as driftless random walks and geometric random walks

$$B_t = B_{t-1} + \nu_t, \tag{4.4}$$

$$\alpha_t = \alpha_{t-1} + \zeta_t, \tag{4.5}$$

$$log\sigma_t = log\sigma_{t-1} + \eta_t, \tag{4.6}$$

where $\nu_t \sim N(0, Q)$, $\zeta_t \sim N(0, S)$ and $\eta_t \sim N(0, W)$. According to Primiceri (2005), the assumptions of random walk processes hold as long as (4.4)-(4.6) are in place for a finite period of time. Also, the hypothesis of random walk enables to focus on permanent shifts and reducing the number of parameters in the estimation process. Q, S and W are positive definite matrices with S assumed to be block diagonal, with blocks corresponding to parameters belonging to separate equations.

Similarly to Primiceri (2005), we employ the same priors with the first 40 observations used to calibrate the prior's distributions. On this subsample, the mean and the variance of B_0 and A_0 are estimated by OLS along with its four times variance in a time invariant VAR. Likewise, for log_{σ_0} the mean is the logarithm of the OLS point estimates of the standard errors and the variance is arbitrarily chosen to be an identity matrix. The Q, W, S matrices follow inverse-Wishart distributions with different degrees of freedom. For Q the degrees of freedom are set to 40, equal to the size of the subsample, as a tighter prior is necessary to avoid implausible behaviours of time varying coefficients. The degrees of freedom are set to 4 for W, 2 and 3 for the two blocks of S, such that the inverse Wishart prior has a finite mean and variance. In summary,

$$
\begin{aligned}
B_0 &\sim N(\widehat{B}_{OLS}, 4 \cdot V(\widehat{B}_{OLS})) \\
A_0 &\sim N(\widehat{A}_{OLS}, 4 \cdot V(\widehat{A}_{OLS})) \\
log\sigma_0 &\sim N(log\widehat{\sigma}_{OLS}, I_n) \\
Q &\sim IW(k_Q^2 \cdot 40 \cdot V(\widehat{B}_{OLS}), 40) \\
W &\sim IW(k_w^2 \cdot 4 \cdot I_n, 4) \\
S_1 &\sim IW(k_S^2 \cdot 2 \cdot V(\widehat{A}_{1,OLS}), 2) \\
S_2 &\sim IW(k_S^2 \cdot 3 \cdot V(\widehat{A}_{2,OLS}), 3)
\end{aligned}
$$

where S_1 and S_2 are the two blocks of S, while $\widehat{A}_{1,OLS}$ and $\widehat{A}_{2,OLS}$ stand for the two corresponding blocks of $\widehat{A}_{OLS}$. The hyperparameters k_Q, k_S and k_W are set equal to 0.01, 0.1 and 0.01, respectively. This way, the priors are not flat, but diffuse and uninformative.

We estimate the model in a two-step principal component approach as it is common in the literature (Stock and Watson (2005), Korobilis (2013), Molteni and Pappa (2017)). First, we estimate the factors F_t by using standard principal components from the data matrix X_t. Then, we include the first principal components in the model and use Bayesian methods to estimate the time varying parameters. Specifically, both parameters and hyperparameters in A_t, B_t, Ω_t, Q, W, S are estimated via the Gibbs sample with conditional posteriors as in Primiceri (2005).

4.2.2 Identification of Monetary policy shocks

We follow Bernanke et al. (2005) to identify monetary policy shocks in recursive order by setting observed factors right after the unobserved ones. Therefore, we order the estimated factors before ECB's Main Refinancing Operations (MRO) rate. This ordering assumes that the estimated factors do not respond to a potential monetary policy shock within the same period. However, we do not need to impose that assumption on all the information variables (here, X_t). Instead, we divide the information dataset in two categories, named: "slow-moving" and "fast-moving". For the first case, variables such as output, unemployment or price indexes will not react to an unexpected monetary shock on impact. "Fast-moving" , on the contrary, respond contemporaneously to policy shocks. Information on the transformation of the series is provided in Appendix. Then, the direct dependence of principal components on the MRO (R_t) rate must be removed. We estimate through a multiple regression the coefficient of R_t

$$\hat{PC}_t = b_c \hat{PC}_t^s + b_r R_t + e_t \tag{4.7}$$

where $\hat{PC}_t^s$ are the principal components extracted from the subset of "slow-moving" variables. $\hat{F}_t$ is finally constusted as $\hat{PC}_t - b_r R_t$.

4.3 Data

We conduct the analysis of monetary transmission for 12 EA countries: Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal and Spain. As is common in the literature, we term the group of Austria, Belgium, Finland, France, Germany and the Netherlands the "core countries" of EA, while we call the group of Greece, Italy, Ireland, Spain and Portugal the "periphery countries". The data sample ranges from 2000:01 to 2019:12 in monthly frequency. Therefore, our sample starts a year after the official start of the EA with the ECB acting as single central bank for the monetary union.

We employ 289 macroeconomic and financial time series which can be described in the following categories: industrial production; unemployment rates; consumer and producer[1] prices; exchange rates; interest rates; share stock prices; historic volatilities; monetary aggregates; price indexes; and miscellaneous. Data are obtained from Deutsche Bundesbank, Eurostat, European Central Bank(ECB) and Organisation for Economic Cooperation and Development(OECD). Series are seasonally adjusted where needed by us-

[1] Data on all components of producer prices are not available for the case of Ireland for the entire sample and therefore are excluded.

ing the X-13 R package, developed by the United States Census Bureau. Stationarity of the series is needed before conducting factor analysis. Therefore, data are transformed by first differences or first differences of logs. Following Barigozzi et al. (2014), Boivin et al. (2008), Laine (2020) and Potjagailo (2017) we obtain first differences of logs for all series except interest rates and unemployment rates. First differences are taken for unemployment rates and consumer/industrial confidence indicators. Then, series are tested for unit roots by the Augmented Dickey-Fuller test. Finally, all series are standardized with zero mean and unit variance. The full list of the series and their transformations can be found in Appendix B.

4.4 Impulse responses in and across Euro Area countries

We start our analysis by examining the impact of a monetary policy to the Euro Area aggregate and across the country members by using a Time Varying Parameter - FAVAR model. Instead of using Bai and Ng (2002) information criterion, as it is common in factor analysis literature, we follow Bernanke et al. (2005), Molteni and Pappa (2017) and Stock and Watson (2002) that the number of factors for the TVP-FAVAR model should be decided through the examination of the sensitivity of the results compared to an alternative number of factors. Particularly, overestimation of the number of factors can reduce the efficiency of the model, whereas, underestimation may lead to inconsistency due to loss of important dynamics. Therefore, we use the first three out of the 239 estimated factors as they explain 12.73%, 8.19% and 5.71% of the variance (almost 27%), respectively. We also include 2 lags as it is common by the TVP-(FA)VAR literature (Korobilis (2013), Molteni and Pappa (2017), Prüser and Schlösser (2020)). The estimation of the Gibbs sampling algorithm is implemented with 10,000 iterations, discarding the first 2000. [2]

[2]Implementation of 20,000 iterations gives similar results.

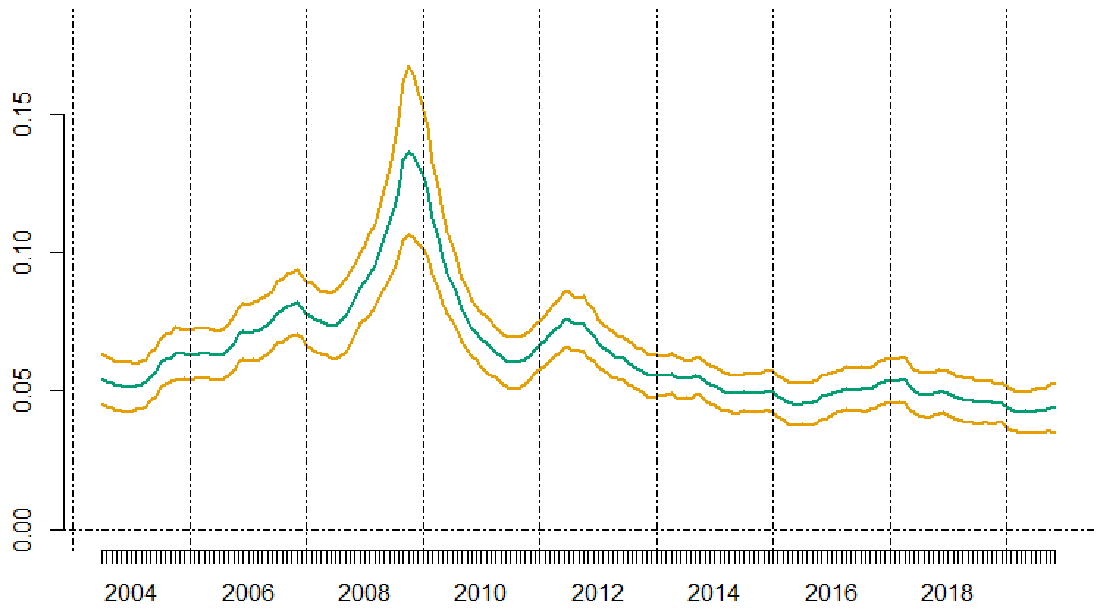


Figure 4.1: Posterior mean of the residuals of the monetary policy shocks along with the 16th and 84th percentiles.

Figure 4.1 shows the posterior mean along with the 16th and 84th percentiles of the time varying standard deviation of the monetary policy shocks. The plot suggests time variation for the volatility of the policy rate in several time periods. First, volatility increases in 2008 reaching a peak (0.147 points) right before the end of the year. In the next years, volatility is fluctuating in lower levels (0.05-0.1) between 2011-2013, returning to its initial levels afterwards. All dates of significant variations are not surprising as they are linked with the financial and European debt crisis starting in 2008 until 2015, whereas some countries already recovered since 2011.

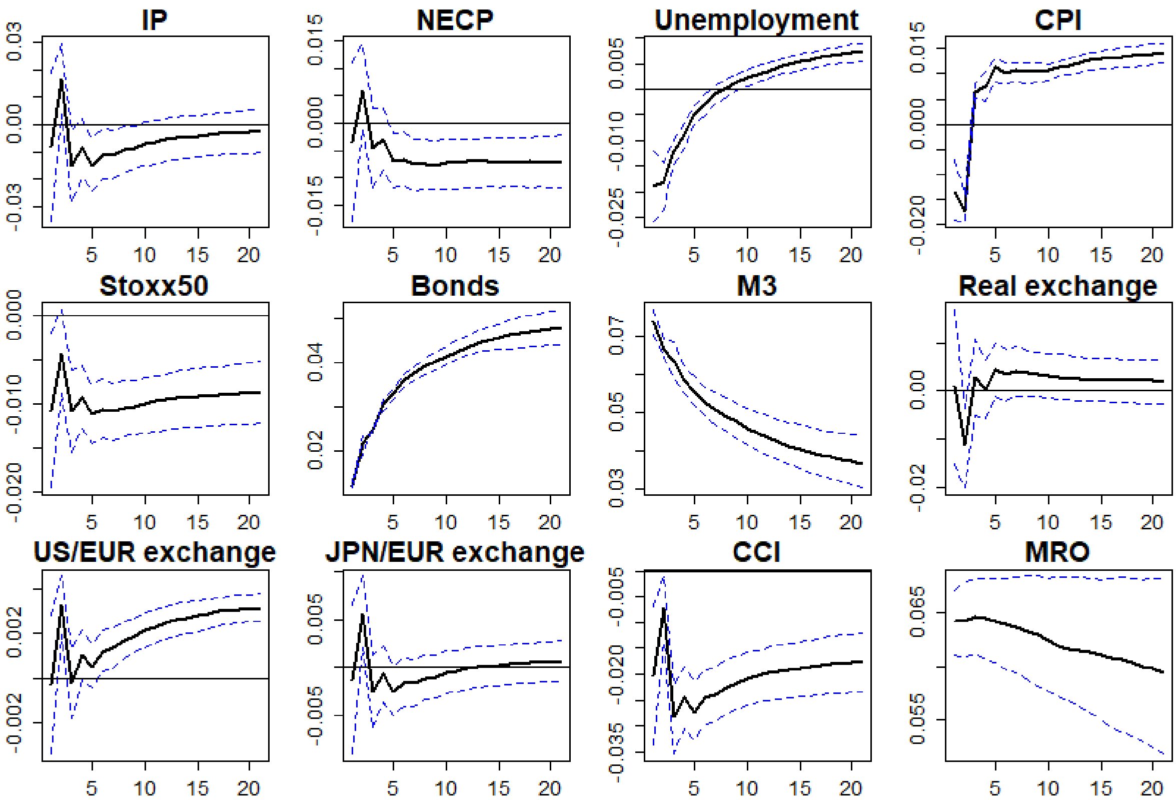


Figure 4.2: Impulse responses to a monetary policy shock on Euro area aggregates. Blue dotted lines represent the 68% confidence intervals.

Figure 4.2 shows the impulse responses of a selection of variables representing the Euro area aggregate to one standard deviation of the error term in the MRO interest rate. The selection covers different sections of economy that play crucial role to the transmission of monetary policies. Specifically, Industrial Production (IP), Non Energy Commodity Price index (NECP), Unemployment rates, Consumer Price Index (CPI), Stoxx50 Index (Stoxx50), Long-term government bonds (Bonds), M3, Real effective exchange rate (Real exchange), US/EUR and JPN/EUR exchange rates, Consumer Confidence Index as CCI and MRO. The bold lines correspond to the median impulse responses and the dotted lines represent the 68%(blue) confidence intervals. The forecast horizon is set to 21 months.

Industrial production and commodity price index react significantly with a delay on a contractionary monetary policy shock. Unemployment, after an initial decline, increase within a year. Furthermore, consumer price inflation falls on impact and then increases to positive levels within a semester after the shock. Concerning the financial variables, stoxx50 declines with a short delay, bonds appreciate and M3 increases on impact up to 0.07 base points, then gradually dies out. Real exchange and JPN/EUR exchange rates are not affected by the monetary shock as they are not statistically significant. US/EUR rates are appreciating six months after the shock. Finally, Consumer Confidence Index's response to an increase in MR0 rate is negative. Overall, the results are consistent with the literature in monetary transmission policies using (TVP)FAVAR models (Bernanke et al. (2005), Galariotis et al. (2018), Korobilis (2013), Molteni and Pappa (2017), Potjagailo (2017), Primiceri (2005)).

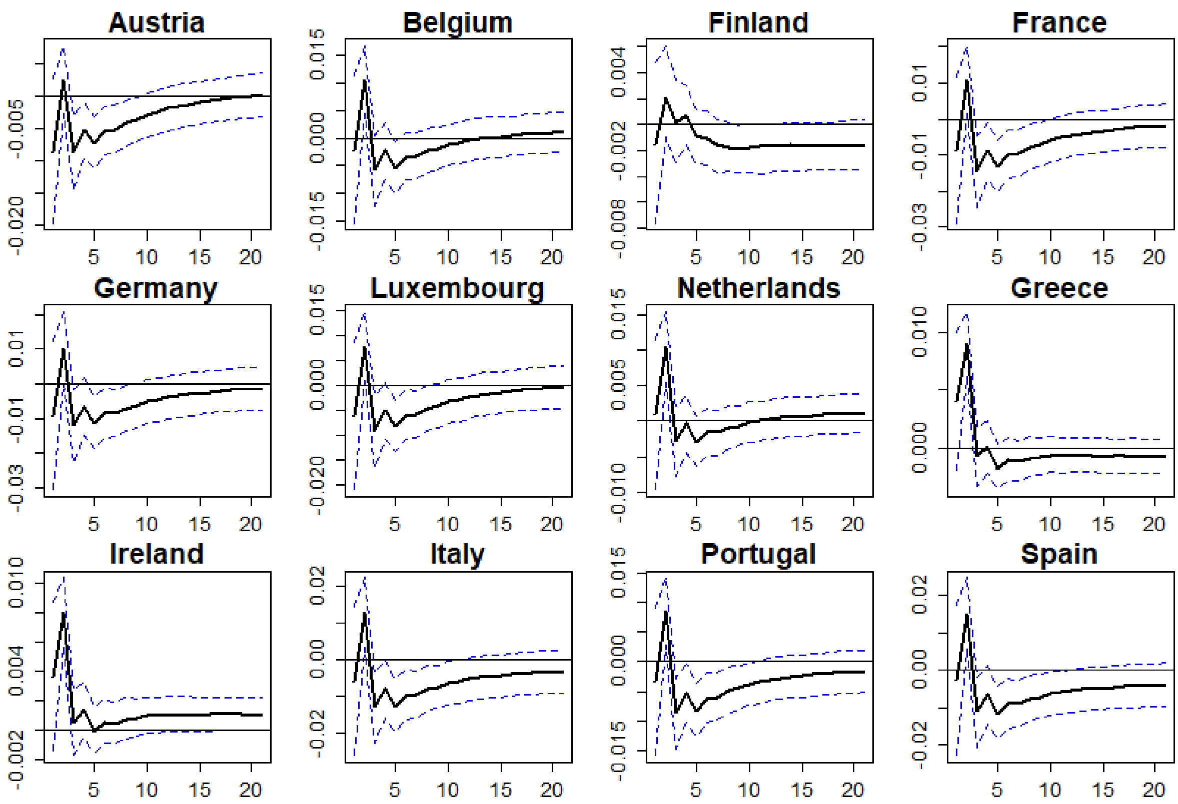


Figure 4.3: Impulse responses to a monetary policy shock on industrial production growth rates. Blue dotted lines represent the 68% confidence bands.

Next, we examine the monetary transmission on 12 Eurozone members for several macroeconomic variables. Figure 4.3 presents the impulse responses of industrial production growth rates, along with the 68% confidence intervals. Industrial production decreases with a short delay almost for all countries. Here, Finland is not statistically significant. Netherlands, Greece and Ireland present a positive effect during the second month after the shock and then return to zero level. Overall, the results show a symmetric spillover experience for the case of industrial production growth in both country groups, similar to the EA19 average.

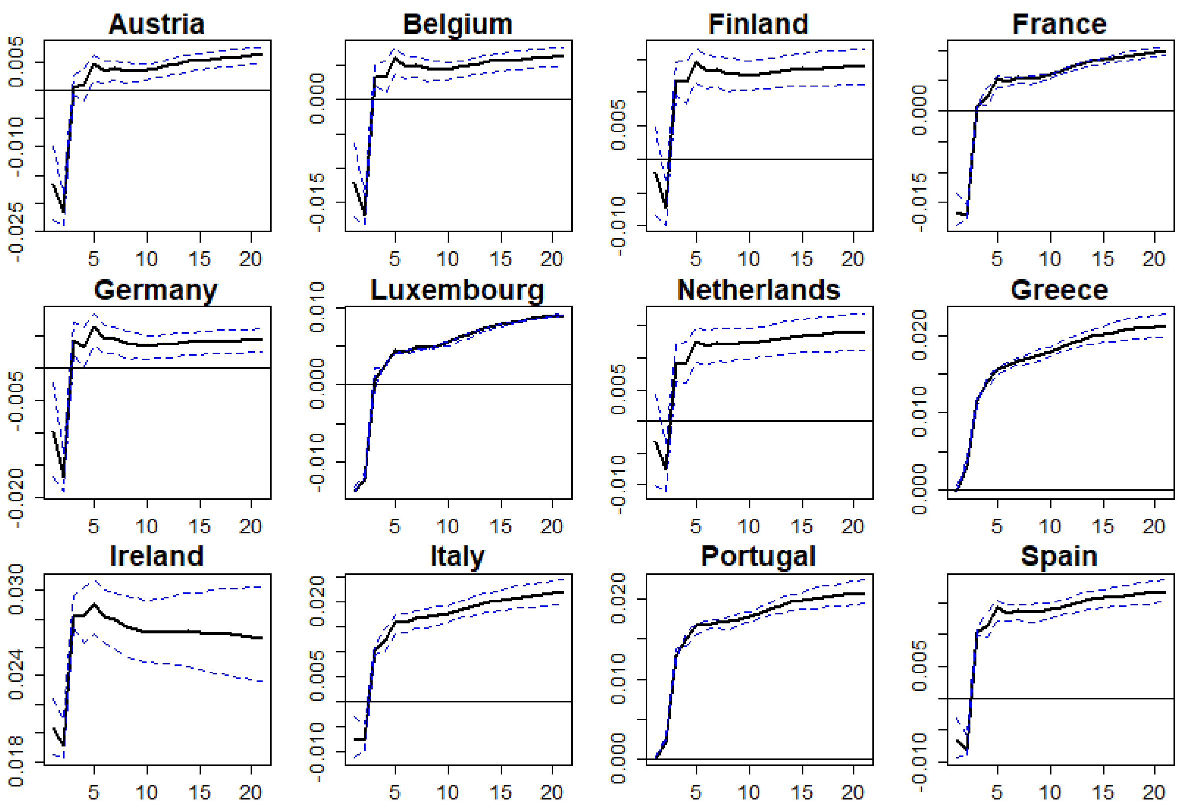


Figure 4.4: Impulse responses to a monetary policy shock on consumer price indexes. Blue dotted lines represent the 68% confidence bands.

Figure 4.4 shows the impulse responses of CPI to each EA country. Here, the spillover effect does not appear as symmetric as in the case of industrial production. For the case of core countries, consumer price inflation rates fall significantly on impact, then increasing and reaching a maximum after five months. France and Luxembourg, instead, present a more persistent behavior as they continue to increase in the long-term. Periphery countries present a positive shock to an unexpected increase of MRO. Specifically, Greece, Ireland and Portugal are increasing on impact, whereas Italy and Spain are initially decreasing and then follow the same direction as the other countries. Overall, the size and the direction of the monetary shock differ among and inside the group-countries. Netherlands and Finland experience the shortest decline at 0.005 base points whereas

Austria,Belgium and Germany experience a stronger one between 0.015 to 0.02 base points. In periphery group, Ireland reaches 0.03 base points five months later whereas the rest of the group countries range in lower levels. Overall, the results show an heterogeneity in prices across Euro Area. Periphery countries do not follow the responses of core countries and the EA19. This could be a result of price rigidities and lack of competition (Barigozzi et al. (2014)).

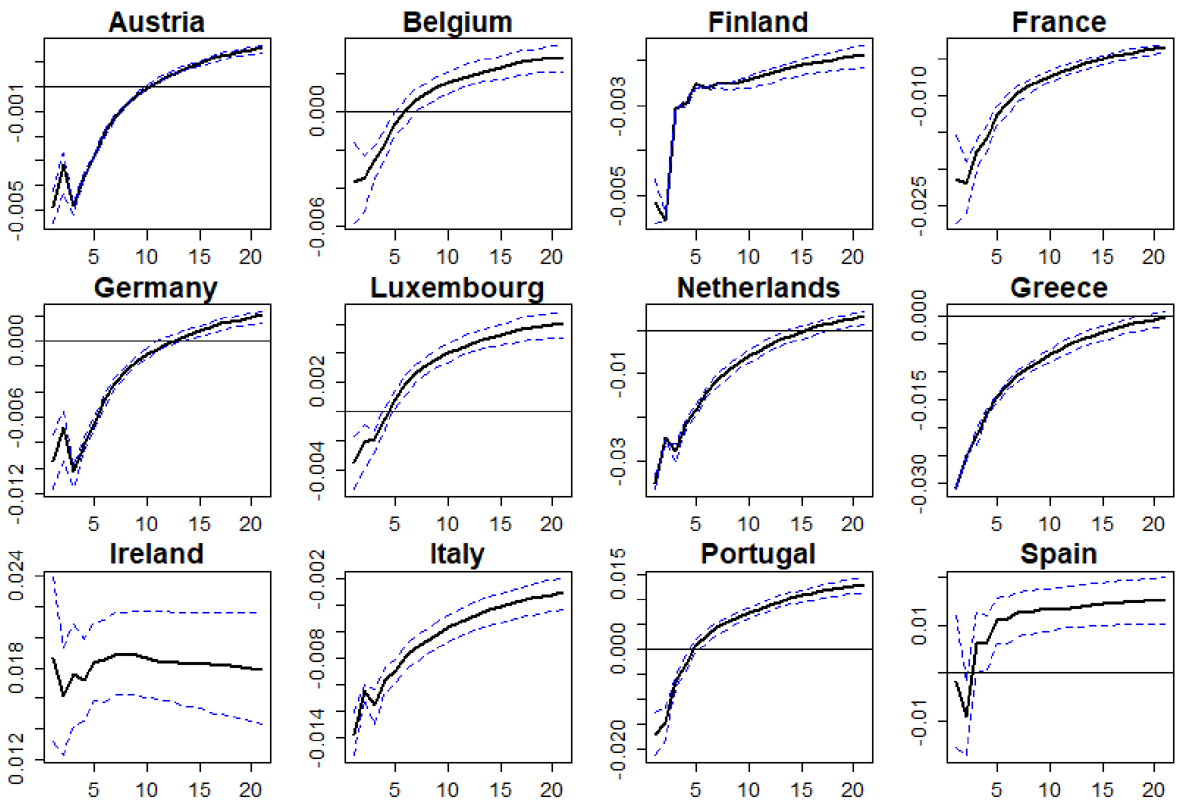


Figure 4.5: Impulse responses to a monetary policy shock on unemployment rates. Blue dotted lines represent the 68% confidence intervals.

Unemployment rates in figure 4.5 decline significantly for both core and periphery countries on impact and finally increase. As it is excepted by the economic theory, the magnitude of the shock differs across the groups. In core countries, the response increases right after the shock towards the positive levels. Austria, Belgium and Luxembourg surpass the baseline within the same year after the shock. For Germany and Netherlands the effect is more persistent and becomes positive after a year and a half. For Finland and France the shock gradually dies out. For the case of periphery countries, Portugal and

Spain display a similar behavior as the the core countries while Ireland's unemployment increases and stays steady for the whole forecast horizon. Greece and Italy decrease on impact and as Finland and France return to the baseline. The spillover effects on unemployment rate in EA appear to move on the same direction with different size impact and durations which can be explained by the different labor regimes each country follows (Altavilla and Ciccarelli (2009)).

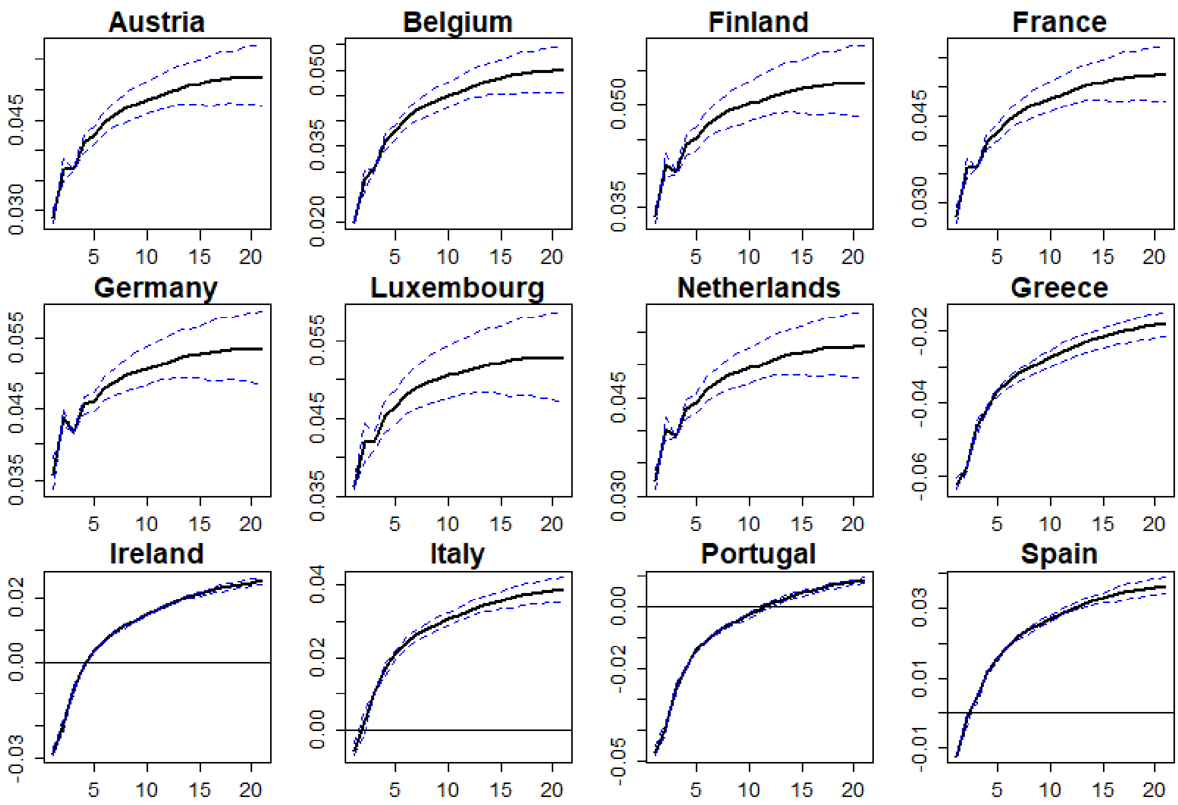


Figure 4.6: Impulse responses to a monetary policy shock on long-term government bonds. Blue dotted lines represent the 68% confidence intervals.

Likewise, long term government bonds in figure 4.6 present asymmetric spillover effects. More specifically, in core countries government bonds increase significantly on impact with the same size, whereas periphery countries react on different directions. Greek government bonds depreciate on impact before slowly returning to zero level. Ireland, Portugal and Spain initially depreciate and finally appreciate. Interestingly, appreciation takes place in different time horizons for each country. Italian government bonds appreciate as core countries and EA19. The responses of government bonds show that the level of economic integration is of great importance for the magnitude of the final effects.

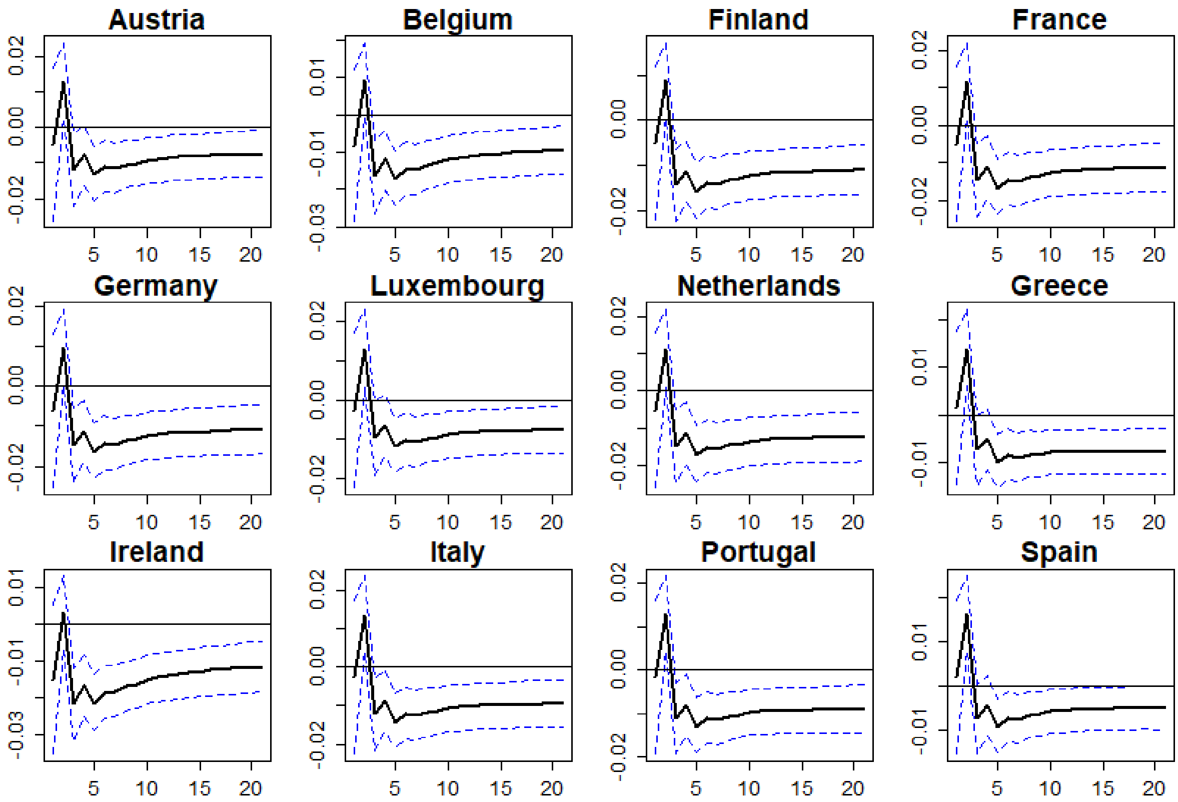


Figure 4.7: Impulse responses to a monetary policy shock on share price indexes. Blue dotted lines represent the 68% confidence intervals.

Finally, share price indexes move on the same negative direction with a delay. The findings are symmetric across all EA countries and in line with the literature that indicates the negative relationship between Monetary policy and stock markets (Lütkepohl and Netšunajev (2018)).

As it is of great importance to assess the monetary policy shocks in different periods of time, we conclude our analysis by examining the monetary policy in three different time periods and also compare our results with a traditional TVP-VAR model for the EA19 average. Policy makers can then investigate if the asymmetries derive from the monetary policy itself or the structure of the economy. Looking at the volatility of the MRO rate in figure 4.1 we first choose October of 2008 as it is the month with the highest volatility as well as a date linked with the beginning of the financial crisis. The other two dates are arbitrarily chosen as four years before the beginning of the crisis and four years after. Thus, we can have a better comparison between the pre-crisis period and the crisis/post-crisis period. Those are October of 2004 and October of 2012.

We first employ a four variable TVP-VAR model. Here, the variables in y_t are industial production growth rates (IP), inflation rates (CPI), MRO rate and unemployment rates. Unemployment rates are placed last as a monetary shock affects labour market on impact. The number of lags is chosen again to 2.

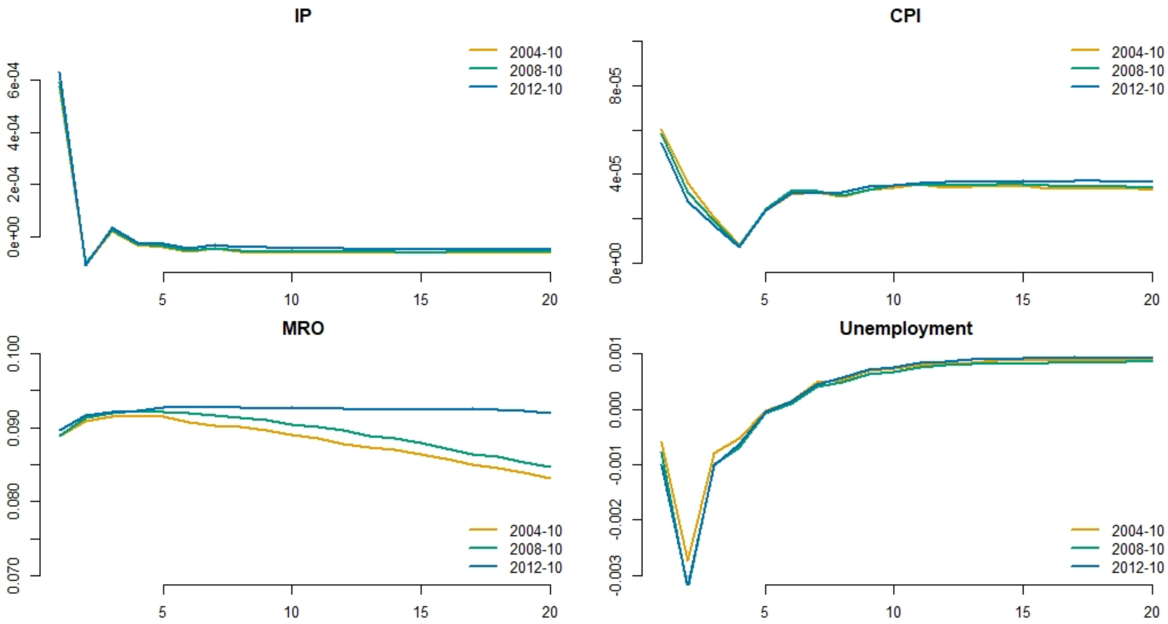


Figure 4.8: Impulse responses to a monetary policy shock on industrial production, inflation, interest rates and unemployment from the TVP-VAR model for periods 2004.10, 2008.10 and 2012.10.

In Figure 4.8 the impulse response functions from the TVP-VAR model for all four included variables at the three different time periods are presented. First, industrial production declines with delay and the estimated responses do not vary much over time. Next, inflation exhibits "price puzzle" as it increases on impact, contrary to economic theory, and stays in positive levels without strong evidence of time variation. On contrary, MRO's shock is positive and more persistent during the post crisis period. Last, unemployment declines on impact and then finally increases for the selected time periods.

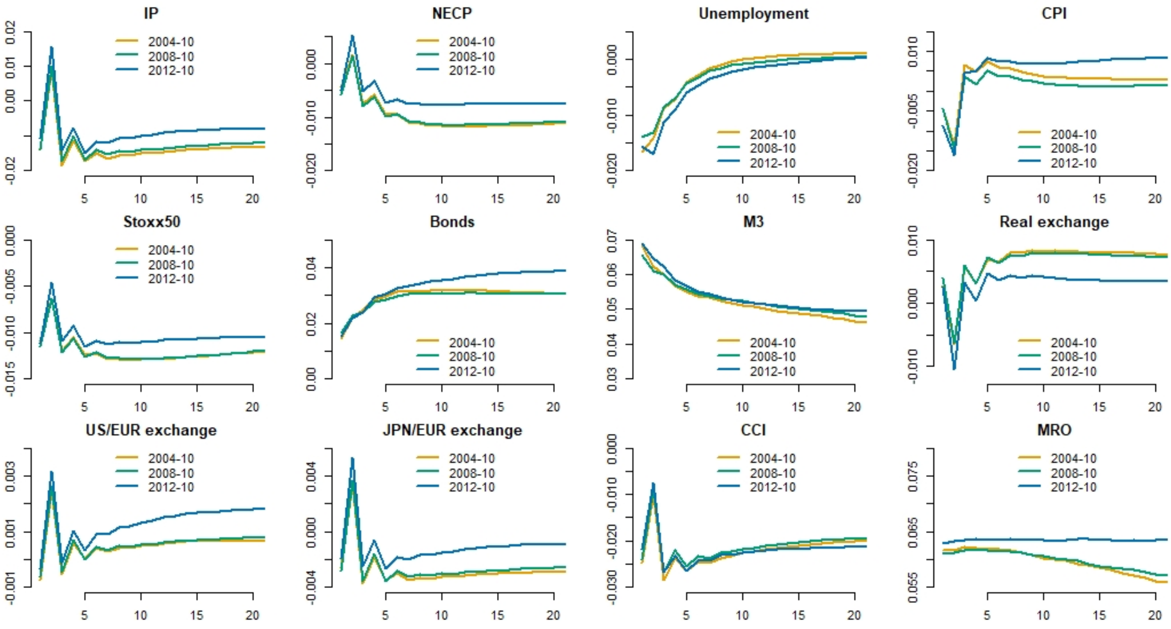


Figure 4.9: Impulse responses to a monetary policy shock on Euro area aggregates for periods 2004.10, 2008.10 and 2012.10.

We conclude by examining the major economic variables in figure 4.9 for the Euro area aggregates for the same dates. The results suggest that the majority of the variables under examination reacted similarly given the different time periods. Comparing the TVP-VAR with the TVP-FAVAR model, we observe by including more information about prices, industrial production and their components that the "price puzzle" is reduced as the prices in the TVP-FAVAR model initially decrease. A significant outcome of the results is the evidence of time variation for the case of the TVP-FAVAR model. For instance, industrial production, commodity prices, stoxx50, real and JPN/EUR exchange rates react more moderately during the post crisis period, when the MRO rate is in lower levels. Unemployment, inflation, long-term government bonds and US/EUR exchange rates seem to be affected from the crisis period and react more persistently compared to the responses of 2004 and 2008. Last, M3 and consumer's confidence indicator do not vary significantly during the selected time periods.

4.5 Conclusions

In this paper, we empirically examine the monetary transmission between the EA countries. To do so, we investigate the effects of a contractionary monetary policy shock on the dynamics of several real and nominal economic variables such as industrial production, unemployment, government bonds, producer and consumer price inflation, real exchange rates, etc. both on aggregate and country levels. We employ a TVP-FAVAR model following Bernanke et al. (2005), Primiceri (2005) and Molteni and Pappa (2017) which allows the coefficients to vary over time independently of political or economic regimes. The analysis covers the full EA period from January, 2000, to December, 2019 for 12 EA countries consisting of both "core" and "periphery" countries. Studying Euro Area as a whole, we find statistical significant responses in most cases after a potential shock. However, on country-wise level, we find evidence of asymmetries in the case of consumer price inflation rates, unemployment rates and long-term government bonds. Finally, we examine the Euro Area under three different time periods for two time varying parameter (FA)VAR models. We show that a TVP-FAVAR model offers a data rich environment that can reduce the "price puzzle" effect and provide more information on the reactions of a monetary shock. The results suggest several implications for macroeconomic policies within the EA: First, that the monetary transmission is affected by complex interrelations between trade effects, domestic shocks, economic structures and macroeconomic policy. Second, differences in the responses after the financial crisis of 2008 are detected which could be explained by the low level of interest rates, uncertainty in consumption and investments or the banking system.

Appendix B

In the following table several abbreviations have been used. Specifically, EA corresponds to the 12 Euro Area countries (Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, Spain), EA19 to the current Euro Area of 19 countries, SA and NSA mean that the time series are seasonally adjusted or nonseanonally adjusted, respectively. S or F refer to the asumption that the variable is either slow- or fast-moving. Last, OECD and ECB refer to Organisation for Economic Co-operation and Development and European Central Bank, respectively.

Table 4.1: Data transformation

Description	Countries	Transformation	Slow/fast	Source
Industrial production, total(2015=100, SA)	EA, EA19	Log-difference	S	OECD
Industrial Prouction, manufacturing(2015=100, SA)	EA, EA 19	Log-difference	S	Eurostat
Industrial Production, capital goods(2015=100, SA)	EA, EA 19	Log-difference	S	Eurostat
Industrial Production, durable consumer goods(2015=100, SA)	EA, EA 19	Log-difference	S	Eurostat
Industrial Production, intermediate goods(2015=100, SA)	EA,EA 19	Log-difference	S	Eurostat
Industrial Production, non-durable goods(2015=100, SA)	EA, EA 19	Log-difference	S	Eurostat
Industrial Production, energy(2015=100, SA)	EA, EA 19	Log-difference	S	Eurostat
Unemployment rate(SA)	EA, EA 19	1st differences	S	Eurostat
Producer Price Index, total(domastic, 2015=100, NSA)	EA, EA 19	Log-difference	S	Eurostat
Producer Price Index, manufacturing(domestic, 2015=100, NSA)	EA, EA 19	Log-difference	S	Eurostat
Producer Price Index, capital goods(domestic, 2015=100, NSA)	EA , EA 19	Log-difference	S	Eurostat
Producer Price Index, durable consumer goods(domestic, 2015=100, NSA)	EA, EA 19	Log-difference	S	Eurostat
Producer Price Index, intermediate goods(domestic, 2015=100, NSA)	EA, EA 19	Log-difference	S	Eurostat
Producer Price Index, non-durable goods(domestic, 2015=100, NSA)	EA, EA 19	Log-difference	S	Eurostat
Producer Price Index, energy(domestic, 2015=100, NSA)	EA, EA 19	Log-difference	S	Eurostat
Consumer Price Index, total(2015=100, NSA)	EA, EA 19,USA,JPN	Log-difference	S	OECD
Consumer Price Index, food(2015=100, NSA)	EA, EA 19	Log-difference	S	OECD
Consumer Price Index, energy(2015=100, NSA)	EA, EA 19	Log-difference	S	OECD
Crude Oil Price (NA)	Worldwide	Log-difference	S	World Bank
Commodity Price Index, non energy (NA)	EA 19	Log-difference	S	ECB
Money market 3-month rates (nominal, NA)	EA 19,USA,JPN	No transformation	F	Eurostat
ECB Main refinancing Repo rate	EA 19	No transformation	F	Deutsche Bundesbank
Long-term interest rates (10 years, NA)	EA 19,USA,JPN	No transformation	F	Eurostat
M1, M2, M3 (percentage change)	EA 19	No transformation	F	ECB
Real effective exchange rate (27 trading parterns, 2010=100)	EA countries EA 19	Log-difference	F	Eurostat
Euro exchange rates	UK,USA,JPN	Log-difference	F	Eurostat
Share price index	EA, USA, JPN	Log-difference	F	OECD
EUROSTOXX50	EA	Log-difference	F	Datastreams
Stock price historic volatility	EA,USA,JPN	No transformation	F	Datastreams
Consumer/Industrial Confidence indicators	EA 19	1st differences	F	OECD

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